Money Illusion and Housing Frenzies*

Markus K. Brunnermeier†
Department of Economics
Princeton University

Christian Julliard‡
Department of Economics
London School of Economics

July 25, 2007

Abstract

A reduction in inflation can fuel run-ups in housing prices if people suffer from money illusion. For example, investors who decide whether to rent or buy a house by simply comparing monthly rent and mortgage payments do not take into account the fact that inflation lowers future real mortgage costs. We decompose the price-rent ratio into a rational component – meant to capture the "proxy effect" and risk premia – and an implied mispricing. We find that inflation and *nominal* interest rates explain a large share of the time-series variation of the mispricing, and that the tilt effect is very unlikely to rationalize this finding.

Keywords: Housing, Real Estate, Inflation, Inflation Illusion, Mortgages, Behavioral Finance
JEL classification: G12, R2.

*We benefited from helpful comments from an anonymous referee, Yakov Amihud, Patrick Bolton, Smita Brunnermeier, John Campbell, James Choi, Albina Danilova, Aureo de Paula, Emir Emiray, Will Goetzmann, Kevin Lansing, Chris Mayer, Alex Michaelides, Stefan Nagel, Martin Oehmke, Maureen O'Hara (the editor), Filippos Papakonstantinou, Lasse Pedersen, Adriano Rampini, Matt Richardson, Bob Shiller, Matt Spiegel, Jeremy Stein, Demosthenes Tambakis, Haibin Zhu and seminar and conference participants at the Bank of England, Cambridge-Princeton conference, CSEF-IGIER symposium, Duke-UNC Asset Pricing Conference, 2006 Econometric Society Winter Meetings in Boston, Federal Reserve Bank of Philadelphia, Harvard University, IIES Stockholm, London Business School, London School of Economics, NBER Behavioral Meetings, Oxford University, Queen-Mary University, RFS Bubble Conference in Indiana, University of Copenhagen, University of Salerno, Wharton School, University of Wisconsin-Madison Real Estate Research conference, and Yale Conference on Behavioral Economics. We also thank the BIS for providing part of the housing data used in this analysis. Brunnermeier acknowledges financial support from the National Science Foundation and the Alfred P. Sloan Foundation.

†Princeton University, Department of Economics, Bendheim Center for Finance, Princeton, NJ 08544-1021, NBER and CEPR, e-mail: markus@princeton.edu, http://www.princeton.edu/∼markus

‡Department of Economics, London School of Economics, Houghton Street, London WC2A 2AE, United Kingdom, and CEPR, e-mail: C.Julliard@lse.ac.uk, http://personal.lse.ac.uk/julliard/

1 Introduction

Housing prices have reached unprecedented heights in recent years. Sharp run-ups followed by busts are a common feature of the time-series of housing prices. Figure 1 illustrates different real housing price indices and shows that this phenomenon has been observed in several countries.

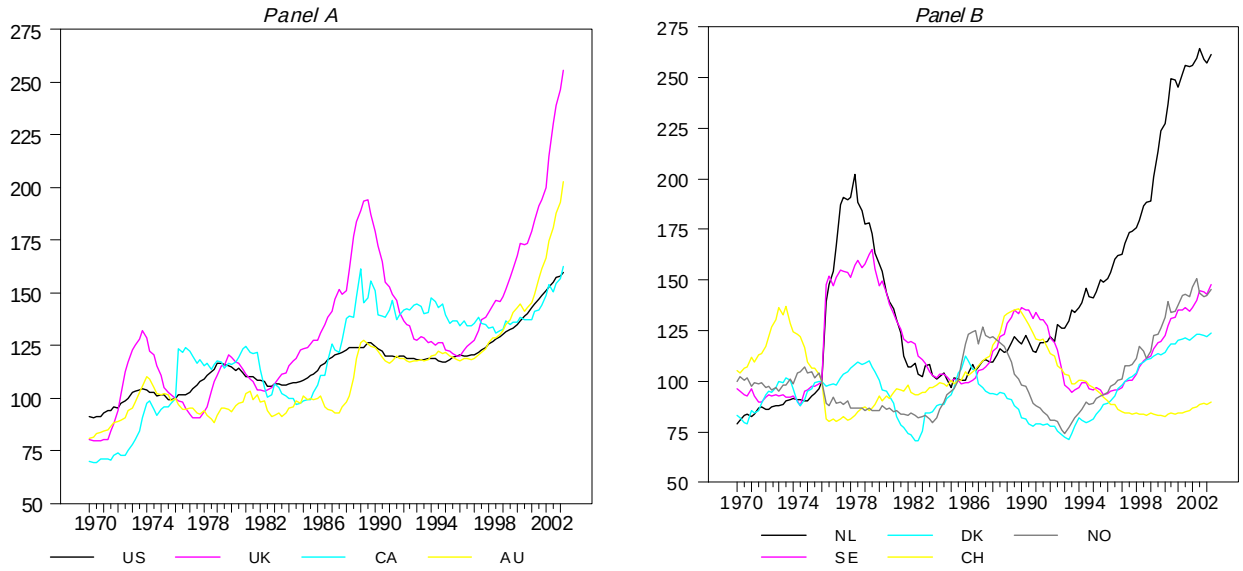

Figure 1: Residential property (real) price indices for a group of Anglo-Saxon countries (Panel A) and for Scandinavian countries and other European countries (Panel B). Base period is 1976:Q1.

Shiller (2005) documents similar patterns for other countries and cities over shorter samples. Moreover, Case and Shiller (1989, 1990) document that housing price changes are predictable and suggest that this might be due to inefficiency in the housing market. There are several potential reasons for this market inefficiency – one of them being money illusion, the inability to properly distinguish changes in nominal values due to changes in real fundamentals from changes merely due to inflation. The housing market is particularly well suited to study money illusion, since frictions, e.g., short-sale constraints, make it difficult for professional investors to arbitrage possible mispricings away.

In this paper we identify an empirical proxy for the mispricing in the housing market and show that it is largely explained by movements in inflation. Inflation matters and it matters in a particular way. Our analysis shows that a reduction in inflation can generate substantial increases in housing prices in a setting in which agents are prone

to money illusion. For example, people who simply base the decision of whether to rent or buy a house on a comparison between monthly rent and monthly payment of a fixed nominal interest rate mortgage suffer from money illusion. They mistakenly assume that real and nominal interest rates move in lockstep. Hence, they wrongly attribute a decrease in inflation to a decline in the real interest rate and consequently underestimate the real cost of future mortgage payments. Therefore, they cause an upward pressure on housing prices when inflation declines.

To identify whether the link between housing price movements and inflation is due to money illusion, we first have to isolate the rational components of price changes that are due to movements in fundamentals, such as land and construction costs, housing quality, property taxes, and demographics [Mankiw and Weil (1989)].[1] We do so in two stages. First, we focus on the price-rent ratio to insulate our analysis from fundamental movements that affect housing prices and rents symmetrically. Even though renting and buying a house are not perfect substitutes, the price-rent ratio implicitly controls for movements in the underlying service flow. Second, we try to isolate rational channels through which inflation could influence the price-rent ratio. Several authors including Fama (1981) have claimed that the negative relationship between inflation and the price of real assets (like stocks) might be due to a "proxy effect": high inflation and/or high inflation expectations are a bad signal about future economic conditions. Moreover, higher inflation might make the economy more risky or agents more risk averse, generating a risk premium that is correlated with inflation. Poterba (1984) stresses a rational channel that implies the opposite effect of inflation on house prices: an increase in inflation reduces the after-tax real user cost of housing, potentially driving up housing demand. We use a Campbell and Shiller (1988) decomposition that takes into account housing-specific risk factors, (such as probability of moving interacted with cross-sectional variation of house prices), to decompose the price-rent ratio into rational components (expected future returns on housing investment and rent growth rates) and a mispricing component. After controlling for rational channels, we find that inflation has substantial explanatory power for the sharp run-ups and downturns of the housing market.

Figure 2 depicts the time series of the (estimated) mispricing component of the price-rent ratio in the U.K. housing market and its fitted values obtained using inflation as the only explanatory variable. The first thing to notice is that the mispricing shows sharp and persistent run-ups during the sample period. Moreover, the fitted series closely tracks the mispricing.

[1] These variables alone are generally not able to capture the sharp run-ups in housing prices. It has become common in the empirical literature to add cubic 'frenzy' terms in the housing price regressions [see Hendry (1984) and Muellbauer and Murphy (1997)] and the rational expectations hypothesis has been rejected by the data [Clayton (1996)].

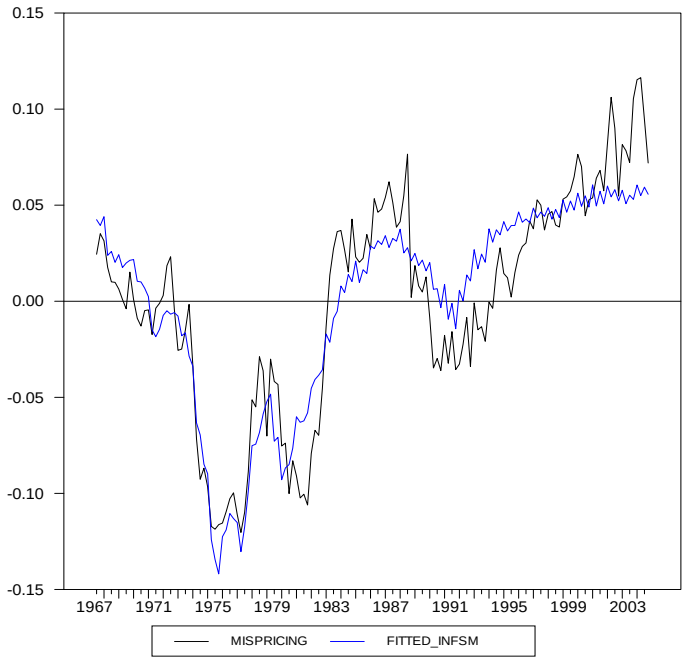

Figure 2: Mispricing and fitted series based on U.K. inflation

The close link between inflation and housing prices could be due to a departure from rationality and/or financing frictions. First, as argued by Modigliani and Cohn (1979), if agents suffer from money illusion, their valuation of an asset will be inversely related to the overall level of inflation in the economy.[2] A special form of money illusion arises if home owners are averse to realizing nominal losses. Second, in an inflationary environment, the nominal payments on a fixed-payment mortgage are higher by a factor that is roughly proportional to the reciprocal of the nominal interest rate. This causes the real financing cost to shift towards the early periods of the mortgage, therefore causing a potential reduction in housing demand and prices. This is the so-called *tilt effect* of inflation [see Lessard and Modigliani (1975); Tucker (1975)]. Nevertheless, why the tilt effect should matter cannot be fully explained in a rational setting since financial instruments that are immune to changes in inflation, like the price level adjusted mortgage (PLAM) or the graduate payment mortgage (GPM), have been available to house buyers since at least the 1970s. Most importantly, in Section 4.1 we perform a series of tests that make it seem very unlikely that the tilt

[2]This explanation of housing price run-ups would also be in line with the finding of McCarthy and Peach (2004) that the sharp run-up in the U.S. housing market since the late 1990s can be largely explained by taking into account the contemporaneous reduction of *nominal* mortgage costs.

effect is the driving force of the empirical link between inflation and housing prices. Third, if fixed interest rate mortgages are not portable, individuals who have bought a house and have locked in a low nominal interest rate might be less willing to sell their current house to buy a better one when nominal interest rates are higher. Hence, an increase in inflation that raises the nominal interest rate might depress the price of better-quality residential properties. On the other hand, a reduction in inflation and nominal interest rates would free current home owners from this "lock-in" effect. We provide evidence that the "lock-in" effect is not driving our results. Further, we show that housing supply elasticity is heterogeneous across U.S. states due to differences in population density. We demonstrate that given this heterogeneity in supply elasticity, money illusion can lead to heterogeneous regional price dynamics as observed in the data [e.g., Gyourko, Mayer, and Sinai (2006)].

The balance of the paper is organized as follows. The next section reviews the related literature on money illusion, market frictions and speculative trading. Section 3 formally analyzes the link between inflation and housing prices using the U.K. housing market as a case study.[3] In particular, Subsection 3.1 provides a first assessment of the empirical link between price-rent ratio and inflation. Subsection 3.2 decomposes the price-rent ratio isolating the rational channels from an estimated mispricing and shows that the mispricing is largely explained by changes in the rate of inflation. Section 4 argues that market frictions – like the tilt effect (Subsection 4.1) and lock-in effect (Subsection 4.2) – are unlikely to be the cause of the link between inflation and mispricing on the housing market. Section 5 confirms the main empirical results using U.S. data and studies the heterogeneity of housing supply elasticity across the U.S. A final section concludes and a full description of the data sources, methodological details, and additional robustness checks are provided in the Appendix.

2 Related Literature

2.1 Money Illusion and Psychological Biases

> "An economic theorist can, of course, commit no greater crime than to assume money illusion," Tobin (1972).
>
> "In fact, I am persuadable – indeed, pretty much persuaded – that money illusion is a fact of life," Blinder (2000).

In this section, we sketch the links to the existing literature. In particular, we review previous definitions of money illusion, relate it to the psychology literature, and

[3] We first focus on the U.K. market since the better quality of the housing data, the longer sample period in housing prices (1966:Q2–2004:Q4) and inflation-linked bonds (1982:Q1-2004:Q4), the availability of PLAM and GPM mortgage schemes, and the fact that most U.K. mortgages are portable, allow for sharper and more robust inference.

summarize the empirical evidence on the effect of money illusion on the stock market.

Definition of Money Illusion. Fisher (1928, p. 4) defines money illusion as "the failure to perceive that the dollar, or any other unit of money, expands or shrinks in value."[4] Patinkin (1965, p. 22) refers to money illusion as any deviation from decision making in purely real terms: "An individual will be said to be suffering from such an illusion if his excess-demand functions for commodities do not depend [...] solely on relative prices and real wealth..." Leontief (1936) is more formal in his definition by arguing that there is no money illusion if demand and supply functions are homogeneous of degree zero in all nominal prices.

Related Psychological Biases. Money illusion is closely related to other psychological judgement and decision biases. In a perfect world, money is a veil and only real prices matter. Individuals face the same situation after doubling all nominal prices and wages. Nevertheless, psychological biases might not allow individuals to see through this veil.

The *framing effect* states that alternative representations (framing) of the same decision problem can lead to substantially different behavior [Tversky and Kahneman (1981)]. Shafir, Diamond, and Tversky (1997) document that agents' preferences depend to a large degree on whether the problem is phrased in real terms or nominal terms. This framing effect has implications for both time preferences and risk attitudes. For example, if the problem is phrased in nominal terms, agents prefer the nominally less risky option to the alternative, which is less risky in real terms. That is, they avoid nominal risk rather than real risk. On the other hand, if the problem is stated in real terms, their preference ranking reverses. The degree to which individuals ignore real terms depends on the relative *saliency* of the nominal versus real frame.

Anchoring is a special form of the framing effect. It refers to the phenomenon that people tend to be unduly influenced by arbitrary quantities when presented with a decision problem. This is the case even when the quantity is clearly uninformative. For example, the nominal purchasing price of a house can serve as an anchor for a reference price even when the real price can be easily derived.[5] Genesove and Mayer (2001) document that investors are reluctant to realize nominal losses.

While individuals understand well that inflation increases the prices of goods they buy, they often overlook inflation effects that work through *indirect channels* (e.g., general equilibrium effects). For example, Shiller (1997a) documents survey evidence

[4]Most authors use the terms "money illusion" and "inflation illusion" interchangeably. Sometimes the latter is also used to refer to a situation where households ignore changes in inflation.

[5]Fisher (1928) provides several interesting examples of inflation illusion due to anchoring. For example, on pages 6-7, he writes about a conversation he had with a German shop woman during the German hyperinflation period in the 1920s: "That shirt I sold you will cost me just as much to replace as I am charging you [...] But I have made a profit on that shirt because I bought it for less."

that the public does not think that nominal wages and inflation comove over the long-run. Shiller (1997b) provides evidence that less than a third of the respondents in his survey study would have expected their nominal income to be higher if the U.S. had experienced higher inflation over the last five years. The impact of inflation on wages is more indirect. Inflation increases the nominal profits of the firm, therefore ceteris paribus it will increase nominal wages. Similarly, the reduction in mortgage rates due to a decline in expected future inflation expectations is direct, while the fact that it will also lower future nominal income is indirect. This inattention to indirect effects can be related to two well-known psychological judgement biases: mental accounting and cognitive dissonance. *Mental accounting* [Thaler (1980)] is a close cousin of narrow framing and refers to the phenomenon that people keep track of gains and losses in different mental accounts. By doing so, they overlook the links between them. In our case, they ignore the fact that higher inflation affects the interest rate of the mortgage and the labor income growth rate in a symmetric way. *Cognitive dissonance* and the *self attribution bias* might be another reason why individuals do not realize that inflation increases future nominal income: people have a tendency to attribute increases in nominal income to their own achievements than simply to higher inflation.[6]

2.2 Inflation and the Stock Market

Several studies document a *negative* correlation between nominal stock returns and inflation – realized and expected [e.g., Lintner (1975); Fama and Schwert (1977); Gultekin (1983)] and unexpected [Amihud (1996)]. This appears puzzling since the Fisher-relation implies that nominal rates should move one-for-one with *expected* inflation. One interpretation of these findings is that inflation proxies for future economic conditions: higher inflation is associated with a grim economic outlook [e.g., Fama (1981)]. On the other hand, it has been argued that the negative correlation might be due to money illusion. Modigliani and Cohn (1979) claim that prices significantly depart from fundamentals since investors make two inflation-induced judgement errors: (i) they tend to capitalize equity earnings at the nominal rate rather than the real rate and (ii) they fail to realize that firms' corporate liabilities depreciate in real terms. Hence, stock prices are too low during high inflation periods. There are many papers that empirically document the impact of money illusion on stock market prices, often referred to as the "Modigliani-Cohn" hypothesis. Ritter and Warr (2002) document that the value-price ratio is positively correlated with inflation and that this effect is more pronounced for leveraged firms. Using Campbell and Shiller's (1988) dynamic log-linear valuation method and a subjective proxy for the equity risk premium, Campbell and Vuolteenaho (2004) show in the time-series that a large part of the mispricing

[6] Shiller (1997a) also noted that "Not a single respondent volunteered anywhere on the questionnaire that he or she benefited from inflation. [...] There was little mention of the fact that inflation redistributes income from creditors to debtors."

in the dividend-price ratio can be explained by inflation illusion.[7] Cohen, Polk, and Vuolteenaho (2005) focus on the cross-sectional implications of money illusion on asset returns and find supportive evidence for the "Modigliani-Cohn" hypothesis. It is worth emphasizing that proxy effect and money illusion are not mutually exclusive.

On the other hand, Boudoukh and Richardson (1993) find that at low frequency nominal market returns are *positively* correlated with inflation consistently with the Fisher relation. This finding is not inconsistent with money illusion: even though investors suffering from money illusion underestimate the nominal earnings growth of companies after an increase in inflation, they should realize their mistake once the actual nominal earnings are announced.

Basak and Yan (2005) show, within a dynamic asset pricing model, that even though the utility cost of money illusion (and hence the incentive to monitor real values) is small, its effect on equilibrium asset prices can be substantial. In the same spirit, Fehr and Tyran (2001) show that (under strategic complementarity) even if only a small fraction of individuals suffer from money illusion, the aggregate effect can be large.

To the best of our knowledge, we are the first to empirically assess the link between money illusion and housing prices. We find strong support in favor of money illusion and weak supportive evidence for the proxy effect.

It should be emphasized that stock and housing markets differ both in their structure and their composition. While the residential housing market is dominated by individual households, institutional investors play a major role in the stock market. Further, trading frictions, most notably short-sale constraints, severely limit arbitrage in the residential housing market.

2.3 Borrowing Constraint and Speculation

Tilt Effect. Lessard and Modigliani (1975) and Tucker (1975) show that under nominal fixed payment and fixed interest rate mortgages, inflation shifts the real burden of mortgage payments towards the earlier years of the financing contract. In the presence of borrowing constraints, this limits the size of the mortgages agents can obtain. This tilt effect could lead to a reduction in housing demand. Kearl (1979) and Follain (1982) find an empirical link between inflation and housing prices and argue that liquidity constraints could rationalize their finding. Wheaton (1985) questions this simple argument in a life-cycle model and shows that several restrictive assumptions are needed for this to be the case.

Speculative Trading and Short-Sale Constraints. In the presence of money illusion and short-sale constraints, the potential disagreement between rational and irrational agents can also lead to housing frenzies.

[7] Additional evidence on the time-series link between market returns and inflation can be found in Asness (2000, 2003) and Sharpe (2002).

Piazzesi and Schneider (2007) show in a static setting that house prices are highest whenever the disagreement about the inflation level is high.[8] Harrison and Kreps (1978) show in a dynamic setting that speculative behavior can arise if agents have different opinions (i.e., non-common priors). Said differently, even if they could share all the available information, they would still disagree about the likelihood of outcomes. Scheinkman and Xiong (2003) put this model in a continuous-time setting and show that transaction costs dampen the amount of speculative trading, but only have limited impact on the size of the bubble. Models of this type rely on the presence of short-sale constraints – which is a natural constraint in the housing market – to preempt the ability of rational agents to correct the mispricing. Other factors that limit arbitrage include noise-trader risk [DeLong, Shleifer, Summers, and Waldmann (1990)] and synchronization risk [Abreu and Brunnermeier (2003)].

Note also that collateral and downpayment constraints – as analyzed in Stein (1995), Bernanke and Gertler (1989), Bernanke, Gertler, and Gilchrist (1996), Kiyotaki and Moore (1997), Ortalo-Magné and Rady (2006) – combined with money illusion would lead to an amplification of the negative effect of inflation on housing prices.

3 Housing Prices and Inflation

We focus on the link between inflation and the price-rent ratio. In principle, an agent could either buy or rent a house to receive the same service flow. However, renting and buying a house are not perfect substitutes since households might derive extra utility from owning a house (e.g., ability to customize the interior, pride of ownership). Moreover, properties for rent might on average be different from properties for sale.[9] Nevertheless, long-run movement in the rent level should capture long-run movements in the service flow. Furthermore, changes in construction cost, demographic changes, and changes in housing quality should at least in the long-run affect housing prices and rent symmetrically. As a consequence, in studying mispricing in the housing market, we focus on the price-rent ratio. Gallin (2004) finds that housing prices and rents are cointegrated and that the price-rent ratio is a good predictor of future price and rent changes. Compared to the price-income ratio, the price-rent ratio has the advantage of being less likely to increase dramatically due to changes in fundamentals (e.g., in demography or property taxes). Moreover, Gallin (2003) empirically rejects the hypothesis of cointegration between prices and income using panel-data tests for cointegration, which have been shown to be more powerful than the time-series analog.

[8] For a discussion of Piazzesi and Schneider (2007), see Brunnermeier (2007).

[9] The house price index reflects all types of dwellings while rents tend to overweight smaller and lower quality dwellings. Given that high-quality houses fluctuate more over the business cycle, the data might show a spurious link between inflation, nominal interest rates and the price-rent ratio if inflation and/or nominal interest rates had a clear business cycle pattern. We address this concern formally in Section 3.2.2 and show that this does not affect our main findings.

This implies that the commonly used error correction representation of prices and income would lead to erroneous frequentist inference. Finally, studying the price-rent ratio is also analogous to the commonly used price-dividend ratio approach to analyze the mispricing in the stock market.

In this section, we show first that a simple non-linear function of the nominal interest rate is a proxy for the valuation of the price-rent ratio by an agent prone to money illusion. Empirically, we first document the correlation between nominal variables and future price-rent ratios. To gain further understanding of this empirical link, we then decompose the price-rent ratio into a rational component and an implied mispricing and study its comovements with inflation. In this section, we conduct our empirical analysis focusing on U.K. data because the longer sample period (1966:Q2–2004:Q4) and the better quality of the data allow us to obtain sharper and more robust inference.

3.1 Housing Prices and Money Illusion - A First-Cut

In a dynamic optimization setting, the equilibrium real price an agent is willing to pay for the house, P_t, should be equal to the present discounted value of future real rents, $\{L_t\}$, and the discounted resale value of the house.

$$P_t = \tilde{E}_t \left[\sum_{\tau=1}^{T-1} m_{t,t+\tau} L_{t+\tau} + m_{t,T} P_T \right],$$

where $m_{t,\tau}$ is the stochastic discount factor between t and $\tau > t$, T is the time of resale, and $\tilde{E}_t$ is the expectations operator given agents' subjective beliefs at time t.

In order to present a first insight into the role of inflation bias, we start by considering a simple setting without uncertainty and with constant real rent, as in Modigliani and Cohn (1979). In this case, as $T \to \infty$, the equilibrium price-rent ratio for an economy with rational agents is:

$$\frac{P_t}{L_t} = E_t \left[\sum_{\tau=1}^{\infty} \frac{1}{(1 + r_{t,t+\tau})^{\tau}} \right] \simeq \frac{1}{r_t}, \tag{1}$$

where $r_{t,t+\tau}$ is the *real* (quarterly) risk-free yield from t to τ, r_t is the real risk-free rate, and we assume that $\lim_{T\to\infty} \left(\frac{1}{1+r_{t,T}}\right)^T P_T = 0$. Equation (1) holds exactly if the real risk-free rate, r_t, is constant.[10]

[10]Note that strictly speaking L_t reflects all payoffs from owning a house. This includes not only the service flow from living in the house but also tax benefits, property tax, etc. For our empirical analysis, we focus only on the main component: the market price of the service from living in the house. The standard user cost approach in real estate economics takes the other components into account as well. The user cost is stated in terms of per dollar of house value. More specifically,

Instead, if the agent suffers from money illusion, the agent treats the (constant) nominal risk-free yield, i, as real. This implies the inflation biased evaluation:

$$\frac{P_t}{L_t} = \tilde{E}_t \left[\sum_{\tau=1}^{\infty} \frac{1}{(1 + r_{t,t+\tau})^{\tau}} \right] \simeq E_t \left[\sum_{\tau=1}^{\infty} \frac{1}{(1 + i_{t,t+\tau})^{\tau}} \right] \simeq \frac{1}{i_t}, \tag{2}$$

where the first approximation ignores the Jensen's inequality term and the second approximation is exact if the nominal interest rate, i_t, is constant.[11] This derivation parallels the one in Modigliani and Cohn (1979) for the stock market. Equations (1) and (2) suggest that $1/i_t$, $1/r_t$ and inflation π_t should be used as alternative regressors to test for money illusion. It is also worth emphasizing that $1/i_t$ is highly non-linear in i_t for low i_t – a fact independently emphasized for the real interest rate by Himmelberg, Mayer, and Sinai (2005).

To take a first look at the empirical link between inflation, nominal interest rates and the price-rent ratio, we explore whether i_t, r_t, π_t, $1/i_t$, and $1/r_t$ have forecasting power for the price-rent ratio. In assessing the forecasting performance of these variables, one faces several econometric issues. First, Ferson, Sarkissian, and Simin (2002) use a simulation exercise to argue that the in-sample regression results may be spurious, and both the R^2 and statistical significance of the regressor are biased upward if both the *expected* part of the regressand and the predictive variable are highly persistent [see also Torous, Valkanov, and Yan (2005)]. Therefore, since P_t/L_t is highly persistent, this could lead to spurious results. Second, in exploring the forecastability of the price-rent ratio, the choice of the control variables is problematic and to some extent arbitrary since the literature on housing prices has suggested numerous predictors. Moreover, Poterba (1991) outlines that the relation between housing prices and forecasting variables often used in the literature has not been stable across sub-samples.

We address both issues jointly. For the first problem, we remove the persistent component of the price-rent ratio by constructing the following forecasting errors

$$\hat{\delta}_{t+1,t+1-\tau} = \begin{cases} P_{t+1}/L_{t+1} - \hat{E}_{t-\tau}\left[P_{t+1}/L_{t+1}\right] & \text{for } \tau > 0 \\ P_{t+1}/L_{t+1} & \text{for } \tau = 0, \end{cases} \tag{3}$$

$u_t = r_t^f + \omega_t - \tau\left(r_t^m + \pi_t + \omega_t\right) + \delta_t - g_{t+1} + \gamma_t$, where r_t^f is the risk-free real interest rate, ω_t the property tax per dollar house value, the third term captures the fact that nominal interest payments and property taxes are deductible from the income tax with marginal tax rate τ, δ_t reflects maintenance costs, and g_{t+1} is the capital gain (loss) per dollar of house value, γ_t is the risk premium. Note that since nominal mortgage interest payments are income tax deductible, inflation lowers user cost and, since the price-rent ratio should be equal to the reciprocal of the user cost, this suggests higher house prices [see Poterba (1984, 1991)]. This is exactly the opposite inflation effect of the one caused by money illusion. A major drawback of the user cost approach is that the house price appreciation is assumed to be exogenous and is not derived from a consistent dynamic equilibrium. In particular, by assuming that the price appreciation follows historical patterns, one implicitly assume "irrational" positive feedback trading phenomena.

[11] Equation (2) makes clear that money illusion matters independently of whether the mortgage contract is a flexible rate or a fixed rate one.

where τ is the forecasting horizon and $\hat{E}_{t-\tau}[P_t/L_t]$ is the (estimated) persistent component of the price-rent ratio and we introduce the convention that for $\tau = 0$, $\hat{\delta}_{t+1,t+1} = P_{t+1}/L_{t+1}$. Second, we estimate $\hat{E}_{t-\tau}[P_t/L_t]$ by fitting a reduced form vector auto regressive model (VAR) for P_t/L_t, the log gross return on housing, $r_{h,t}$, the rent growth rate Δl_t and the log real return on the twenty-year government bonds, r_t (constructed as the nominal rate, i_t, minus quarterly inflation).[12]

Following Campbell and Shiller (1988), for small perturbations around the steady state, the variables included in the VAR should capture most of the relevant information for the price-rent ratio. Indeed, the R^2 of the VAR equation for P_t/L_t is about 99%, which is consistent with previous studies that have outlined the high degree of predictability of housing prices [see, among others, Kearl (1979); Follain (1982); Muellbauer and Murphy (1997)]. This approach for constructing forecast errors, $\hat{\delta}_{t+1,t+1-\tau}$, is parsimonious since it allows us to remove persistency from the dependent variable without assuming a structural model. It is also conservative since the reduced form VAR is likely to over-fit the price-rent ratio. We use quarterly data over the sample period 1966:Q3–2004:Q4. The VAR is estimated with one lag since this is the optimal lag length suggested by both the Bayesian and Akaike information criteria.

Figure 3 summarizes the results about the predictability of the price-rent ratio. The figure plots Newey and West (1987) corrected t-statistics (Panel A) and measures of fit (Panel B) of five univariate regressions of $\hat{\delta}_{t+1,t+1-\tau}$ on r_t, i_t, $1/r_t$, $1/i_t$ and a smoothed moving average of inflation, π_t.[13] (Recall that we introduced the convention that for $\tau = 0$, $\hat{\delta}_{t+1,t+1} = P_{t+1}/L_{t+1}$.) That is, the first point in each of the plotted series corresponds to the regression output of a standard forecasting regression for the price-rent ratio.

Focusing first on $\tau = 0$ – the standard forecasting regression – it is apparent that the real interest rate, r, has no forecasting power for the price-rent ratio with a t-statistic (Panel A) of 0.741 and a R^2 (Panel B) of about 0%. This is consistent with the finding of Muellbauer and Murphy (1997) that the real interest rate has no explanatory power for movements in the real price of residential housing. The sign of the slope coefficient of the nominal interest rate, i, is negative suggesting that an increase in the nominal interest rate reduces the price-rent ratio. The regressor is statistically significant only at the 10% level and explains about 5% of the variation in the price-rent ratio. The

[12] Note that one could alternatively remove the persistent component of the regressors. But doing this, would add an additional layer of uncertainty since our ability of removing the persistent component might change from regressor to regressor. Furthermore, this alternative approach would put too much emphasis on the last innovation of the regressor.

Also note that we reject that the price-rent ratio is non-stationary, consistent with findings in Gallin (2004). As a consequence, we cannot model P_t/L_t as cointegrated with any of the regressors considered.

[13] Note that the measure of inflation we use is the CPI index *without* housing. The smoothing window is of sixteen quarters and we take 0.9 as the smoothing parameter.

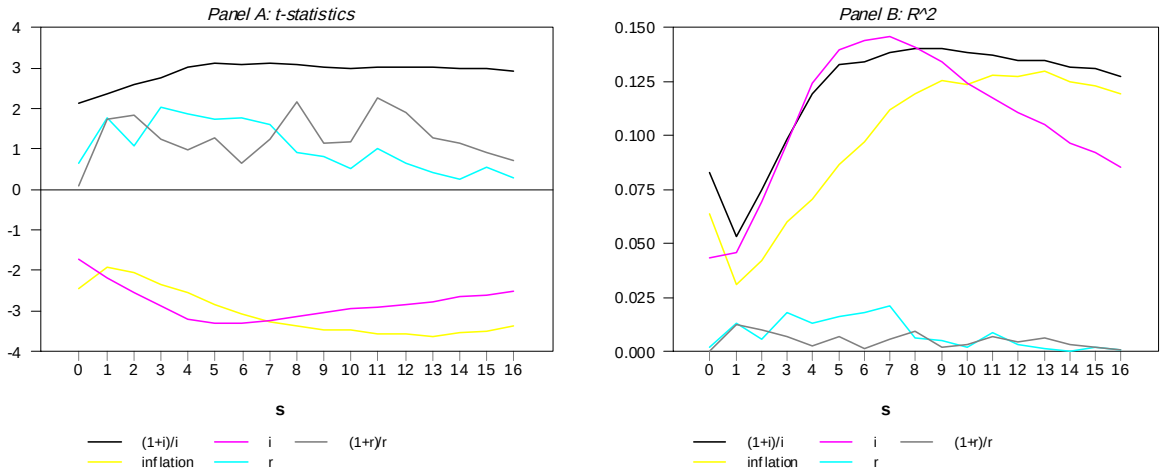

Figure 3: t-statistics and R^2 of univariate regressions of the forecast error $\hat{\delta}_{t+1,t+1-\tau}$ on interest rates and interest rate reciprocals (both nominal and real), as well as inflation.

figure also shows that lagged inflation is a significant predictor of the price-rent ratio and that the estimated slope coefficient has a negative sign, which is consistent with the Modigliani and Cohn (1979) argument that inflation causes a negative mispricing in assets. This is also consistent with the findings of Kearl (1979) and Follain (1982) that housing demand is reduced by greater inflation. The regressor explains about 7% of the time variation in P_t/L_t. From the predictive regression of the price-rent ratio on $1/r_t$ – as suggested by Equation (1) – we learn that this variable is not significant nor has any forecasting power for the future price-rent ratio, reinforcing the conjecture that housing prices do not tend to respond to changes in the real interest rate. However, the reciprocal of the nominal interest rate, $1/i_t$, is highly statistically significant and has a positive sign implying that the price-rent ratio tends to comove with the valuation of agents prone to money illusion. Moreover, this regressor is able to explain about 9% of the time variation in the price-rent ratio. Consistently with money illusion, inflation π_t shows a significant negative correlation with housing prices.

Focusing on $\tau > 0$, we can assess whether the regressors considered have forecasting power for the unexpected component of price-rent changes.[14] It is clear from Figure 3 that the real interest rate (both in terms of r and $1/r$) generally has no explanatory power for the unexpected movements in the price-rent ratio. To the contrary, the nominal interest rate, inflation, and the reciprocal of the nominal interest rate are

[14]Recall that if the results obtained with $\tau = 0$ are due to the persistence of regressors and regressand, we would expect the statistical significance of the regressors to be substantially reduced when considering $\tau > 0$.

statistically significant forecasting variables of unexpected movements in the price-rent ratio, and explain a substantial share of the time series variation of this variable.

For robustness we check our results using the real interest rate implied by the yields on inflation protected ten-year government bonds, instead of using nominal interest rate minus inflation, and using the implied inflation instead of our smoothed inflation. Unfortunately, this data is available only since 1982:Q1. Consistently with the previous results, we find that this measure of the real interest rate also has no explanatory power for the price-rent ratio: the regressor is not statistically significant for any horizon τ and its point estimate changes sign at some horizons. Moreover, using implied inflation instead of smoothed inflation we obtain similar patterns as in Figure 3. The only difference is that implied inflation is not statistically significant at two horizon levels, $\tau = 1$ and 2; this is likely to be due to the fact that we lose sixteen years of quarterly data using implied inflation. Similarly, the real yield spread does not seem to matter. We define the real yield spread as the ten-year real interest rate from inflation protected government bonds minus the three-month government bills reduced by current inflation. Moreover, estimated real interest rate variability and inflation variability are generally not significant predictors of the price-rent ratio, but nevertheless add (very little) explanatory power when considered jointly with inflation. The *nominal* yield spread seems to matter, but this might be spurious since its predictive power goes away when we control for the persistent component of the price-rent ratio. Finally, the default spread, defined as the difference in yield between the Great Britain Corporate Bond Yield and the ten-year government bond, has predictive power.[15] Nevertheless, the default spread does not substantially reduce the statistical significance of our main nominal regressors (π_t, $1/i_t$, and i_t).

Case and Shiller (1989, 1990) find that housing price changes are predictable and argue that this might be at odds with market efficiency. To check whether this potential departure from market efficiency is connected with money illusion, we test whether lagged inflation and the reciprocals of the nominal and real interest rates help to predict the first difference of the price-rent ratio. We find that (i) lagged inflation and nominal interest rates explain 6% to 10% of the time series variation of the changes in the price-rent ratio, (ii) these regressors are statistically significant at levels between 1% and 5%, (iii) the estimated signs are consistent with money illusion, and (iv) the real interest rate does not have any predictive power for changes in the price-rent ratio.

Of course, our results only show that the implicit stochastic discount factor is related to inflation. That is, the forecastability of the price-rent ratio could also be due to predictable changes in the required risk-premium. This would be rational, hence it does not need to be caused by money illusion. We disentangle the role of money illusion in the next subsection.

[15] We use the corporate default spread as a proxy for the credit market condition. Ideally, one would like to use the spread between mortgage rates and government bond yields, but this is not feasible due to data limitations.

3.2 Decomposing the Inflation Effect

Inflation can affect the price-rent ratio for various reasons. In this subsection, we differentiate the rational effects of inflation on the price-rent ratio – through expected future rent growth rates and expected future returns on housing – from the irrational, mispricing effect of inflation.

3.2.1 Methodology

We follow the Campbell and Shiller (1988) methodology, but allow agents to have a subjective probability measure, potentially different from the rational probability measure.

Let P and L respectively be the observed (possibly distorted) price and rental payment of housing. The gross return on housing, R_h, is given by the following accounting identity:

$$R_{h,t+1} := \frac{P_{t+1} + L_{t+1}}{P_t}.$$

Under the assumption that the price-rent ratio is stationary, we can log-linearize the last equation around the steady state to get:

$$r_{h,t+1} = (1-\rho)\, k + \rho\,(p_{t+1} - l_{t+1}) - (p_t - l_t) + \Delta l_{t+1},$$

where $r_{h,t} := \log R_{h,t}$, $p_t := \log P_t$, $l_t := \log L_t$, $\Delta l_t := l_t - l_{t-1}$, $\rho := 1/\left(1 + \exp(\overline{l-p})\right)$, $\overline{l-p}$ is the long-run average rent-price ratio (such that $0 < \rho < 1$), and k is a constant. Rearranging the above equation and iterating forward, the log price-rent ratio can be written (disregarding a constant term) as a linear combination of future rent growth, future returns on housing, and a terminal value, i.e.,

$$p_t - l_t = \sum_{\tau=1}^{T} \rho^{\tau-1} \left(\Delta l_{t+\tau} - r_{h,t+\tau}\right) + \rho^T \left(p_{t+T} - l_{t+T}\right). \tag{4}$$

Moving to excess rent growth rates, $\Delta l^e_{t+\tau} = \Delta l_t - r_t$, and excess returns (risk premia) on housing, $r^e_{h,t} = r_{h,t} - r_t$, where r_t is the real return on the long-term government bond (with maturity of 10 or 20 years) and letting T go to infinity, the price-rent ratio can be expressed as:

$$p_t - l_t = \sum_{\tau=1}^{\infty} \rho^{\tau-1} \left[\Delta l^e_{t+\tau} - r^e_{h,t+\tau}\right] + \lim_{T\to\infty} \rho^T \left(p_{t+T} - l_{t+T}\right). \tag{5}$$

This equality between the observed log price-rent ratio, $p_t - l_t$, and future excess rent growth rate, $\Delta l^e_{t+\tau}$, and risk premia, $r^e_{h,t+\tau}$, holds for any realization of $\left(\left\{\Delta l^e_{t+\tau} - r^e_{h,t+\tau}\right\}_{\tau=1}^{\infty}, p_\infty - l_\infty\right)$, and hence holds in expectation for any probability measure.

ψ–Mispricing Measure. Taking expectations of the observed log price-rent ratio in Equation (5) and assuming that the transversality conditions hold, yields:

$$\begin{aligned} p_t - l_t &= \sum_{\tau=1}^{\infty} \rho^{\tau-1} E_t \left[\Delta l^e_{t+\tau}\right] - \sum_{\tau=1}^{\infty} \rho^{\tau-1} E_t \left[r^e_{h,t+\tau}\right] && (6) \\ &= \sum_{\tau=1}^{\infty} \rho^{\tau-1} \tilde{E}_t \left[\Delta l^e_{t+\tau}\right] - \sum_{\tau=1}^{\infty} \rho^{\tau-1} \tilde{E}_t \left[r^e_{h,t+\tau}\right], && (7) \end{aligned}$$

where E_t is the objective expectation operator conditional on the information available at time t and $\tilde{E}_t$ denotes investors' subjective (and potentially distorted) expectations conditional on the same information set. These equalities hold because both rational and irrational investors (ignoring financial frictions) are indifferent to marginal changes to their investment, since the current price-rent ratio is equal to their expected future rent-growth and risk-premia. In particular, investors who require a high risk-premium, $\tilde{E}_t \left[r^e_{h,t+\tau}\right]$, also expect a high expected future rent growth rate, $\tilde{E}_t \left[\Delta l^e_{t+\tau}\right]$, and so support the observed price-rent ratio.

Note that if there are irrational investors, the observed price-rent ratio $p_t - l_t$ can potentially deviate from the true "fundamental value." In this case, the realized excess returns $r^e_{h,t+\tau}$ and rational investors' equilibrium holdings (and potentially the rent growth rate) are distorted, and hence the required risk premium $E_t \left[r^e_{h,t+\tau}\right]$ changes. It is this change in the equilibrium risk premia guarantees that Equations (6) and (7) hold at the *observed* price level.

Note that irrational investors perceive the risk-premium to be $\tilde{E}_t \left[r^e_{h,t+\tau}\right]$, while their actual risk-premium is only $E_t \left[r^e_{h,t+\tau}\right]$. Adding and subtracting $\sum_{\tau=1}^{\infty} \rho^{\tau-1} E \left[\Delta l^e_{t+\tau}\right]$ from the second equation yields:

$$p_t - l_t = \sum_{\tau=1}^{\infty} \rho^{\tau-1} E \left[\Delta l^e_{t+\tau}\right] - \sum_{\tau=1}^{\infty} \rho^{\tau-1} \tilde{E}_t \left[r^e_{h,t+\tau}\right] + \underbrace{\sum_{\tau=1}^{\infty} \rho^{\tau-1} \left(\tilde{E}_t - E_t\right) \left[\Delta l^e_{t+\tau}\right]}_{\psi_t :=}, \quad (8)$$

where the last term, ψ_t, can equivalently be written as:

$$\psi_t = \sum_{\tau=1}^{\infty} \rho^{\tau-1} \left(\tilde{E}_t - E_t\right) \left[r^e_{h,t+\tau}\right]. \quad (9)$$

We use the convention $\left(\tilde{E}_t - E_t\right)[x] := \tilde{E}_t[x] - E_t[x]$. If subjective and objective expectations were to coincide, ψ_t would be zero.

Within the money illusion hypothesis, we identify ψ_t as a mispricing component. In order to see how this definition of mispricing can capture money illusion, let's consider the following example. As in Modigliani and Cohn (1979), individuals fail to distinguish between nominal and real rates of returns. They mistakenly attribute a decrease

(increase) in inflation π_t to a decrease (increase) in real returns, $r_{h,t}$ – or equivalently ignore that a decrease in inflation also lowers nominal rent growth rate ($\Delta l_t + \pi_t$), i.e., $\tilde{E}_t [\Delta l_{t+\tau}] = E_t [\Delta l_{t+\tau} - \pi_{t+\tau}]$. Therefore, our mispricing measure reduces to:

$$\psi_t = -\sum_{\tau=1}^{\infty} \rho^{\tau-1} E_t [\pi_{t+\tau}]. \tag{10}$$

That is, the mispricing and hence the price-rent ratio are increasing as expected inflation declines. Note that in this particular case money illusion always causes a negative mispricing error. However, if individuals have a reference level of inflation, say $\bar{\pi}$, this is not necessarily true. In this case, the last equation becomes:

$$\psi_t = -\sum_{\tau=1}^{\infty} \rho^{\tau-1} E_t [\pi_{t+\tau} - \bar{\pi}]. \tag{11}$$

Even though the level of mispricing is different with a reference level of inflation, its correlation with inflation is unchanged.

In order to make the theoretical decomposition operational for an empirical analysis, we will model rational expectations through a Vector Autoregression (VAR) approach and subjective expectation of the risk premia via a linear factor representation.

First, to construct the empirical counterpart of ψ_t, we follow Campbell (1991) and compute the objective expectations of rent growth rates using a reduced form VAR. The variables included in the VAR are the log excess return on housing, $r^e_{h,t}$, the log price-rent ratio, $p_t - l_t$, the excess rent growth rate, Δl^e_t, and the exponentially smoothed moving average of inflation, π_t. The VAR is estimated using quarterly data and the chosen lag length is one (both the Bayesian and the Akaike information criteria prefer this lag length for the estimated model). Using Equation (6), we obtain the empirical counterpart of $\sum_{\tau=1}^{\infty} \rho^{\tau-1} E_t \left[r^e_{h,t+\tau}\right]$ by subtracting estimated expected rent growth terms from the log price-rent ratio.

Second, to construct the empirical counterpart of ψ_t, we need a proxy for the unobserved term $\tilde{E}_t \left[r^e_{h,t+\tau}\right]$. We follow Campbell and Vuolteenaho (2004) and assume that $\tilde{E}_t \left[r^e_{h,t+\tau}\right]$ is governed by a set of risk factors $\boldsymbol{\lambda}_t$. Hence, we can write $\sum_{\tau=1}^{\infty} \rho^{\tau-1} \tilde{E}_t \left[r^e_{h,t+\tau}\right] = a + \mathbf{b}_1 \boldsymbol{\lambda}_t + \xi_t$. In order to determine ψ_t, presented in Equation (9) as $\sum_{\tau=1}^{\infty} \rho^{\tau-1} \tilde{E}_t \left[r^e_{h,t+\tau}\right] - \sum_{\tau=1}^{\infty} \rho^{\tau-1} E_t \left[r^e_{h,t+\tau}\right]$, we run an OLS regression of $\sum_{\tau=1}^{\infty} \rho^{\tau-1} E_t \left[r^e_{h,t+\tau}\right]$ on the risk factors $\boldsymbol{\lambda}_t$ as follows:

$$\sum_{\tau=1}^{\infty} \rho^{\tau-1} E_t \left[r^e_{h,t+\tau}\right] = \underbrace{a + \mathbf{b}_1 \boldsymbol{\lambda}_t + \xi_t}_{=\sum_{\tau=1}^{\infty} \rho^{\tau-1} \tilde{E}_t \left[r^e_{h,t+\tau}\right]} - \psi_t. \tag{12}$$

We use different potential risk factors. As suggested in Campbell and Vuolteenaho (2004), we use as a first risk proxy the conditional volatility of an investment that

is long on housing market and short on the ten-year government bonds. That is, we construct $\hat{\psi}_t$ as the OLS residual of the following linear regression:

$$\widehat{\sum_{\tau=1}^{\infty} \rho^{\tau-1} E_t} \left[r^e_{h,t+\tau} \right] = \hat{\alpha} + \sum_{\tau=0}^{8} \hat{b}_\tau \hat{h}_{t-\tau} - \hat{\psi}_t, \tag{13}$$

where the first term is constructed as $\widehat{\sum_{\tau=1}^{\infty} \rho^{\tau-1} E_t} \left[r^e_{h,t+\tau} \right] := (p_t - l_t) - \sum_{\tau=1}^{\infty} \rho^{\tau-1} \hat{E}_t \left[\Delta l^e_{t+\tau} \right]$ with $\hat{E}_t \left[\Delta r^e_{t+\tau} \right]$ being the τ-steps ahead VAR forecasts conditional on the data observed up to time t. The regressors $\hat{h}_{t-\tau}$ include seven lagged GARCH-estimates of the conditional volatility[16] and a lagged VAR forecast of the left-hand side variable. The latter acts as a control in an attempt to remove ξ_t from the residual $\hat{\psi}_t$. By doing so, we take a conservative approach in order not to overestimate the mispricing. We also report results using only seven lagged GARCH-estimates of the conditional volatility. We denote this alternative construction of the mispricing by $\hat{\psi}'_t$.

Note that if a house is never sold, the owner is not exposed to any housing market risk except a potential reduction in borrowing capacity. The risk comes about when someone has to buy or sell a house. This is, for example, the case when an individual has to move between areas with different house price levels. Hence, an interaction between the probability of moving and cross-regional variability of house prices is a natural candidate for a risk factor. We therefore introduce a proxy for this source of risk among the risk factors $\boldsymbol{\lambda}_t$. This is done by adding three additional regressors in Equation (13): the cross-regional price variability across the fourteen main macro regions of the U.K., the total within country migration normalized by total population, and the interaction between these two variables. We denote the corresponding mispricing measure by $\hat{\psi}''_t$. Finally, we also experimented with the canonical Fama-French risk factors.

Some note of caution is appropriate about this decomposition. First, the measure of mispricing ψ_t can depend crucially on the chosen subjective risk factor $\boldsymbol{\lambda}_t$ – which is arbitrary. Second, for the OLS construction in Equation (13) to be correct, $\boldsymbol{\lambda}_t$ should be orthogonal to ψ_t. Third, in deriving our ψ-mispricing, we also assume that irrational investors understand the iterated accounting identity in Equation (4).

ε–Mispricing Measure. To derive the ψ-mispricing, we assumed that the transversality condition holds under both the objective and the subjective measure. We now relax this assumption and allow for explosive paths. Moreover, we avoid having to specify exogenous risk factors, $\boldsymbol{\lambda}$, to identify the implied mispricing due to explosive paths.

We define a new measure of mispricing, ε_t, that under the null hypothesis of rational pricing should be zero or at least orthogonal to proxies for money illusion. This

[16] The fitted model is a ARCH-GARCH(2,2) with an AR(1) component for the mean to take into account the persistence in housing returns.

mispricing captures the difference in expectations about future excess rent growth rates and housing investment risk premia plus $\tilde{E}_t\left[\lim_{T\to\infty}\rho^T\left(p_{t+T}-l_{t+T}\right)\right]$:

$$\varepsilon_t := \sum_{\tau=1}^{\infty}\rho^{\tau-1}\left(\tilde{E}_t - E_t\right)\left[\Delta l^e_{t+\tau} - r^e_{h,t+\tau}\right] + \tilde{E}_t\left[\lim_{T\to\infty}\rho^T\left(p_{t+T}-l_{t+T}\right)\right]. \tag{14}$$

That is, ε_t is the difference between observed log price-rent ratio and the log price-rent ratio that would prevail if (i) all agents were computing expectations under the objective measure and (ii) the transversality condition under the objective measure holds, i.e., $E_t\left[\lim_{T\to\infty}\rho^T\left(p_{t+T}-l_{t+T}\right)\right]=0$.

The ε-mispricing can be expressed as a violation of the transversality condition under the objective measure:

$$p_t - l_t = \sum_{\tau=1}^{\infty}\rho^{\tau-1}E_t\left[\Delta l^e_{t+\tau} - r^e_{h,t+\tau}\right] + \underbrace{E_t\left[\lim_{T\to\infty}\rho^T\left(p_{t+T}-l_{t+T}\right)\right]}_{=\varepsilon_t}.$$

To see this, take subjective expectation of Equation (5) and subtract the above equation from it. Therefore, the ε-mispricing captures bubbles that are due to potentially exploding paths, including the intrinsic bubbles analyzed in Froot and Obstfeld (1991). The price patterns depicted in Figure 1 make it difficult to rule out a priori explosive paths over certain subsamples. That is, imposing the objective transversality condition might be too strong an assumption. An explosive path might occur if, for example, agents fail to understand that all the future realizations of returns and rent growth rates must map into the current price-rent ratio as Equation (4) implies. Note that we assume that all traders have the same subjective measure. If traders have heterogeneous measures and face short-sale constraints [as, for example, in Harrison and Kreps (1978)], ε_t could also be affected by a speculative component.

To see how the ε-mispricing relates to money illusion consider, as we did for the ψ-mispricing, the Modigliani and Cohn (1979) benchmark. In this case, we obtain the same result as in Equation (10) and (11) with ψ_t replaced by ε_t. That is, money illusion implies a negative correlation between the ε-mispricing and π_t, i_t, and $-\log\left(1/i_t\right)$.

To estimate this mispricing, we decompose the observed log price-rent ratio into three components: the implied pricing error, $\hat{\varepsilon}_t$, the discounted expected future rent growth, and the discounted expected future returns as follows:

$$p_t - l_t = \sum_{\tau=1}^{\infty}\rho^{\tau-1}\hat{E}_t\left[\Delta l^e_{t+\tau}\right] - \sum_{\tau=1}^{\infty}\rho^{\tau-1}\hat{E}_t\left[r^e_{h,t+\tau}\right] + \hat{\varepsilon}_t, \tag{15}$$

where $\hat{E}_t$ denotes conditional expectations computed using the estimated VAR described above; that is, $\hat{E}_t\left[\Delta l^e_{t+\tau}\right]$ and $\hat{E}_t\left[\Delta d^e_{t+\tau}\right]$ are the τ-step ahead VAR forecasts conditional on the data observed up to time t.

We identify $\hat{\varepsilon}$ under the null that it should be simply a stationary approximation error. If the null is violated, the ε-mispricing could theoretically follow a martingale process (e.g., in the presence of rational bubbles). To take this possibility into account in the empirical sections below, we do two things. First, in the regression analysis presented in Section 3.2.2, we consider both $\hat{\varepsilon}$ and its first difference as dependent variable. Second, since the potential presence of non-stationary behavior in either the ε mispricing and/or the variables included in our VAR specification could lead to complications in the empirical analysis, in Section 3.2.3 we use a Bayesian approach that is immune from stationarity issues.

In summary, we study the following four mispricing measures. These are three different construction of the ψ-mispricing, where the difference arise due to the different sets of risk factors and control variables used for its construction, and the ε-mispricing that does not require the specification of subjective risk factors.

3.2.2 Empirical Evidence

In this subsection, we focus on the empirical links between mispricing measures and inflation. Our first-cut analysis in Section 3.1 showed that nominal terms covary with price-rent ratio rather than real terms. But this link might be due to rational channels, frictions or money illusion. There are several rational channels through which inflation could affect housing prices. First, if inflation damages the real economy, $\sum_{\tau=1}^{\infty} \rho^{\tau-1} E_t \left[\Delta l^e_{t+\tau}\right]$ should be negatively related with inflation. For example, this could be the case of stagflation caused by a cost-push shock. Second, $\sum_{\tau=1}^{\infty} \rho^{\tau-1} E_t \left[r_{h,t+\tau}\right]$ could tend to rise if inflation makes the economy riskier (or investors more risk averse), therefore driving up the required excess return on housing investment. If any of these were the case, the negative correlation between the price-rent ratio and inflation could simply be the outcome of negative real effects of inflation or of time varying-risk premia on the housing investment.

Most importantly, if there were no inflation illusion, we would expect our mispricing measures to be uncorrelated with π_t, $\log(1/i_t)$, and i_t. Instead, the Modigliani and Cohn (1979) hypothesis of money illusion would predict a negative correlation between our mispricing measures and inflation (and the nominal interest rate), and a positive correlation between the mispricing and $\log(1/i_t)$.

In Table 1, Panel A reports the regression output of the three components of the log price-rent ratio in Equation (8), on the exponentially smoothed moving average of inflation, π_t, the nominal interest rate, i_t, and the log of its reciprocal, $\log(1/i_t)$.

Dependent Variables:	Regressors:					
	π_t		i_t		$\log(1/i_t)$	
	Slope coeff.	R^2	Slope coeff.	R^2	Slope coeff.	R^2
Panel A:						
$\hat{\psi}_t$	-4.09 (13.479)	.83	-6.80 (11.765)	.74	.136 (8.020)	.69
$\sum_{\tau=1}^{\infty} \rho^{\tau-1} \hat{E}_t \Delta l^e_{t+\tau}$	-2.58 (2.390)	.12	-3.96 (1.938)	.09	.093 (2.083)	.12
$-\sum_{\tau=1}^{\infty} \rho^{\tau-1} \tilde{E}_t r^e_{h,t+\tau}$	1.92 (1.066)	.03	3.60 (.931)	.03	$-.050$ (.595)	.02
Panel B:						
$\hat{\psi}'_t$	-6.15 (2.483)	.17	-10.9 (2.668)	.17	.241 (2.823)	.19
$\hat{\psi}''_t$	-2.60 (4.812)	.53	-4.79 (5.898)	.55	.092 (4.825)	.49
$\hat{\varepsilon}_t$	-3.90 (7.946)	.65	-6.30 (6.927)	.55	.129 (5.991)	.52

Table 1: Univariate regressions on inflation, nominal interest rate, and illusion proxy. Newey and West (1987) corrected t-statistics in brackets.

The first row of Panel A in Table 1 reports the univariate regression output of regressing the pricing errors on the proxies that are meant to capture inflation illusion. All the regressors are highly statistically significant and the estimated signs are those we would expect under money illusion: the mispricing of the price-rent ratio tends to rise as inflation and nominal interest rates decrease and $\log(1/i_t)$ rises. Moreover, our proxies for inflation bias are able to explain between 69% and 83% of the time series variation of the mispricing of the price-rent ratio. Fitted values computed using inflation are plotted versus the observed values of $\hat{\psi}_t$ in Figure 2 in the introduction. The figure makes clear that the high explanatory power of inflation is not due to a particular subsample.

Ideally, we would like to regress $\hat{\psi}_t$ on the objective expectation of future inflation. One way to capture variations in expected inflation is to use the series of implied inflation from the inflation protected ten-year government bonds. Using this measure as explanator of $\hat{\psi}_t$ we obtain an R^2 of 51% and a point estimate for the slope coefficient of -5.06 with a t-statistics of 4.864.[17]

The second row of Panel A in Table 1 shows that expected future real rent growth rates seem to be negatively correlated with inflation and nominal interest rate (this last variable is significant only at the 10% level), and positively correlated with $\log(1/i_t)$. Nevertheless, only a small share (between 9% and 12%) of the time variation in expected rent growth is explained by the regressors considered. These results are consistent with

[17]Note that in this case, due to data availability problems, we use a sample starting in 1982:Q1.

a view in which inflation influences the price-rent ratio partially due to the fact that high inflation seems to proxy for a worsening of future economic conditions [e.g., Fama (1981)]. On the other hand, this could simply be the outcome of housing rents being more sticky than the general price level.

The third row of Panel A in Table 1 outlines that there is no significant link between inflation and (subjectively expected) risk premia on the housing investment. The regressors considered are not statistically significant and explain only between 2% and 4% of the time series variation in expected future returns on housing. Moreover, the estimated signs of the regressors imply that inflation is associated with a lower risk premium on housing investment (i.e., in times of high inflation the housing investment is considered to be relatively less risky than investing in long-horizon government bonds). Since we use a before-tax measure of returns on housing, this result could also be due to the fact that an increase in inflation increases the after-tax return on housing [see Poterba (1984)], therefore requiring a lower before-tax risk premium.

The sum of the slope coefficients associated with each of the regressors in Table 1 Panel A is an estimate of the elasticity of the price-rent ratio with respect to that regressor. Our results therefore imply that, on average, a 1% increase in inflation (nominal interest rate) maps into a 4.75% (7.16%) decrease in the price of housing relative to rent, and that the largest contribution to this negative elasticity is given by the effect of inflation (nominal interest rate) on the mispricing

Panel B of Table 1 reports the regression coefficients for alternative measures of mispricings. Recall that: ψ_t' is the mispricing constructed without adding controls in Equation (13)[18]; ψ_t'' is the mispricing constructed adding our "moving risk factors"; and that ε_t is the mispricing constructed without specifying exogenous risk factors and measures the mispricing that maps into a violation of the transversality condition under the objective measure. Note that due to data limitations, the ψ_t'' time series starts only in 1975:Q1, while the time series of the other mispricing measures run from 1966:Q3 to 2004:Q4.

The first thing of interest is to compare the sizes of the mispricing of ψ and ψ'. Figure 4 plots the price-rent ratio, and both ψ-mispricing measures over our sample period.

First, notice that the measures of mispricing in Figure 4 generally have the right pattern of correlation with the price-rent ratio. Second, the ψ-mispricing and ε-mispricing capture a non-negligible fraction of the variation in the price-rent ratio. Third, as argued in the methodological section, the ψ'-mispricing measure seems to attribute too large a fraction of the movements in the price-rent ratio to the mispricing. The ψ''-mispricing measure (not depicted), available over the shorter sample 1975:Q1-2004:Q4, closely tracks the ψ and ε-mispricings.

Next, we analyze the explanatory power of the inflation illusion proxies for the ψ'-

[18] We also tried as alternative risk factors the canonical Fama-French risk factors and obtained similar results as for the covariance of $\hat{\psi}_t'$ and the money illusion proxies.

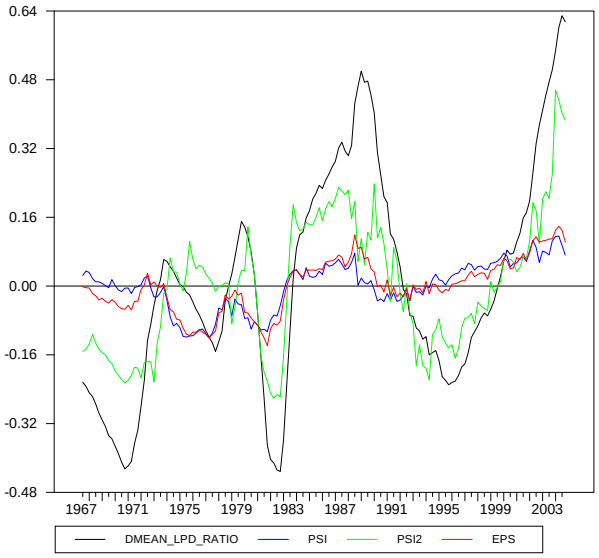

Figure 4: Price-rent ratio and mispricing measures

mispricing and the ε-mispricing. The first row of Panel B of Table 1 shows that both $\hat{\psi}_t'$ and $\hat{\psi}_t''$ – as inflation illusion would imply – covaries negatively (and significantly) with inflation π_t. Similarly, the univariate regressions with nominal interest rate i_t and $\log(1/i_t)$ also deliver significant results consistent with money illusion. Overall, the explanatory power of the inflation illusion proxies is reduced for the ψ'-mispricing. This is not surprising since $\hat{\psi}_t'$ in Figure 4 seems to overstate the time-variation of the mispricing. There is also a small reduction in the measure for the ψ''-mispricing (second row of Panel B), but this might be partially due to the different sample period.

The third row of Panel B of Table 1 reports the regression coefficient of the ε-mispricing on proxies of money illusion. Once again, the signs are consistent with money illusion. Moreover, the estimated elasticities are fairly close to the ones obtained using $\hat{\psi}_t$. Note that theoretically the ε-mispricing could follow a martingale process (e.g., if the price process contains a rational bubble component). Hence, for robustness we also regress the first difference of ε on inflation. The estimated regression coefficient is -4.02 with a t-statistic of 7.459 and an R^2 of 31%.

One worry might be that credit standards might vary over time in response to overall economic conditions, and that this mechanism might generate the link between mispricing and inflation we find in the data. This is potentially important since we have already observed in Section 3.1 that there is a statistically significant link between price-rent movements and the default spread (which is meant to capture the overall economic condition of the credit market). To assess the relevance of the time-variation

of credit market conditions, we regress our mispricing measures on inflation and the default spread jointly. We find that, for all the measures of mispricing, default spread is not statistically significant after controlling for inflation and that the measures of fit do not increase by more than 1%.

Next, the mispricing might be linked to the volatility of inflation more than the level itself. We check this hypothesis by running multivariate regressions of our mispricing measures on inflation and an estimate of conditional inflation volatility.[19] We find that the conditional volatility of inflation has no explanatory power for both mispricing measures after controlling for the level of inflation.

Overall, the results in Table 1 suggest that money illusion can explain a large share of the mispricing in the housing market and that the negative correlation between inflation and the rent-price ratio is mainly due to the effect of money illusion on the mispricing. Nevertheless, our findings could be rationalized by some forms of market frictions. Section 4 addresses this alternative hypotheses formally.

3.2.3 Robustness Analysis

Assessing Uncertainty. To assess the robustness of these results, we next consider the uncertainty due to the fact that we do not directly observe expected rent growth rates and expected future returns on housing, but instead we use the estimated VAR to construct their proxies.

Under a diffuse prior, the posterior distribution of the estimated VAR can be factorized as the product of an inverse Wishart and, conditional on the covariance matrix, a multivariate normal distribution as follows:

$$\begin{aligned} \beta|_{\Sigma} &\sim N\left(\hat{\beta}, \Sigma \otimes (X'X)^{-1}\right) \\ \Sigma^{-1} &\sim \text{Wishart}\left(\left(n\hat{\Sigma}\right)^{-1}, n-m\right), \end{aligned}$$

where β is the vector of slope coefficients in the VAR system, Σ is the covariance matrix of the residuals, the variables with a hat denote the corresponding estimates, X is the matrix of regressors, n is the sample size and m is the number of estimated parameters per equation [see Zellner (1971); Schervish (1995); Bauwens, Lubrano, and Richard (1999)].

This result is exact under normality and the Jeffreys' prior $f(\beta, \Sigma) \propto |\Sigma|^{-(p+1)/2}$ (where p is the number of left-hand side variables), but can also be obtained, under mild regularity conditions, as an asymptotic approximation around the posterior MLE. The Jeffreys' prior formulates the idea of "lack of prejudice" on the space of distribution for the data, and is also flat over the space of the βs and remains flat under reparameterization.

[19] The fitted model is a ARCH-GARCH(1,2) with an AR(1) component for the mean and quarterly dummies to take into account potential seasonality.

The use of this Bayesian approach allows us to draw inference that is robust to the potential presence of non-stationary behavior in both the variables included in our VAR and of our measures of mispricing. The reason being that the likelihood will have an asymptotically Gaussian shape even in the presence of unit roots (Kim (1994)). To summarize the shape of the posterior distribution of the parameters of interest, we compute 10,000 draws from the posterior distribution of the VAR coefficients and, for each draw, we construct expected excess returns, expected rent growth rates and implied mispricing, and use this variables to repeat the regressions reported in the previous section (the procedure is described in detail in Appendix A.2). Table 2 reports the results of this Monte Carlo exercise.

DepVar:	Regressors:					
	π_t		i_t		$\log(1/i_t)$	
	coeff.	R^2	coeff.	R^2	coeff.	R^2
Panel A:						
$\hat{\psi}_t$	-3.10 [-7.79, $-.185$]	.61 [.03, .92]	-5.28 [-12.63, $-.25$]	.57 [.04, .78]	.107 [.01, .25]	.54 [.04, .71]
$\sum_{\tau=1}^{\infty} \rho^{\tau-1} \hat{E}_t \Delta l^e_{t+\tau}$	-2.6 [-11.8, 9.08]	.27 [0, .85]	-4.01 [-18.1, 13.9]	.20 [0, .64]	.095 [$-.303$, .392]	.21 [0, .58]
$-\sum_{\tau=1}^{\infty} \rho^{\tau-1} \tilde{E}_t r^e_{h,t+\tau}$	1.81 [-10.41, 9.61]	.10 [0, .64]	3.44 [-15.34, 15.43]	.09 [0, .59]	$-.048$ [$-.328$, .286]	.07 [0, .44]
Panel B:						
$\hat{\varepsilon}_t$	-3.9 [-11.1, $-.185$]	.64 [.05, .94]	-6.28 [-17.4, $-.68$]	.54 [.05, .75]	.129 [.01, .372]	.52 [.05, .67]

Table 2: Median and 95% confidence intervals for slope coefficients and R^2.

Each row of Table 2 reports the median slope coefficient associated with the regressor, the median R^2, and (in squared brackets) their 95% confidence intervals. The first row of Panel A of Table 2 shows that the relation between inflation illusion and the mispricing of the rent-price ratio is a robust one: inflation and nominal interest rate show a significantly negative correlation with the mispricing, while the inflation-biased valuation shows a significantly positive correlation. Moreover, even though the distribution of the estimated R^2 has a heavy left-tail, there seems to be a very high posterior probability that these variables explain a large share of the time series variation in the mispricing. The second and third rows of Panel A of Table 2 show instead that there is substantial uncertainty about the correlation between inflation, nominal interest rate and expected future returns on housing and expected future rent growth rates. Overall, these results confirm an empirically strong link between nominal values and the mispricing of the housing market, and suggest that this mechanism is the main source of the negative correlation between the price-rent ratio and inflation and the nominal interest rate.

Note that these results are conditional on the estimated risk factor λ_t. The reason being that the uncertainty about λ_t hinges more upon what the risk factor should be than upon how it is estimated. To address this we perform a similar robustness exercise using the ε-mispricing – that does not depend on exogenous risk factors. These results are reported in Panel B of Table 2 and – as in Panel A – are very similar to the ones in Table 1.

Assessing the Role of the Business Cycle. Unlike the price-dividend ratio in the stock market, the observed price-rent ratio is a less precise measure since the housing price index reflects all types of dwellings, while the rent index tends to overweight smaller and lower quality dwellings.

The prices of high-quality houses appreciate at a higher rate during booms, and depreciate more during recessions, than cheaper houses do [see, among others, Poterba (1991) and Earley (1996)]. This might cause the measured price-rent ratio to comove with the business cycle. Hence, if inflation and the nominal interest rate had a clear business cycle pattern, our estimated mispricing measures could show a spurious correlation with these variables.

Figure 8 in Appendix A.3 plots the time series of the U.K. exponentially smoothed quarterly inflation, the return on the twenty-year government bonds, and the Hodrick and Prescott (1997) filtered estimate of the GDP business cycle. The figure shows that there is no strong contemporaneous correlation of inflation and nominal interest rates with the business cycle (the correlation coefficients are $-.16$ and $-.15$, respectively). This suggests that the high degree of explanatory power that inflation and the nominal interest rate have for the housing market mispricing is unlikely to be due to the comovement of these variables with the business cycle. In Appendix A.3 we address this issue formally, and we find that the inclusion of the business cycle in the OLS regressions for the mispricing measures (i) does not drive out the statistical significance of π_t, i_t and $\log(1/i_t)$, (ii) does not significantly change the point estimates of the elasticities of the mispricing reported in Table 1, (iii) does not significantly increase our ability to explain the time variation in the mispricing, and (iv) that the business cycle alone has very little (in the case of $\hat{\psi}_t$ and $\hat{\varepsilon}_t$) or no (in the case of $\hat{\psi}'_t$) explanatory power for the mispricing measures.

4 Market Frictions

The previous section documents a strong link between mispricings in the housing market and inflation, and suggests money illusion as the driving mechanism of this relationship. However, a potential alternative explanation of this finding is the presence of housing market frictions. In this section, we formally investigate this competing hypothesis.

4.1 Tilt Effect

Our empirical results are consistent with money illusion. Nevertheless, we could be capturing the tilt effect of inflation, which potentially generates a negative relationship between inflation and housing prices. The tilt effect refers to a particular form of liquidity constraint. It is best understood by comparing the real repayment profiles of a mortgage with and without inflation. Suppose agents can only enter fixed nominal repayment mortgages. The real repayment profiles of such a contract are depicted in Figure 5 for a zero and a positive inflation environment.

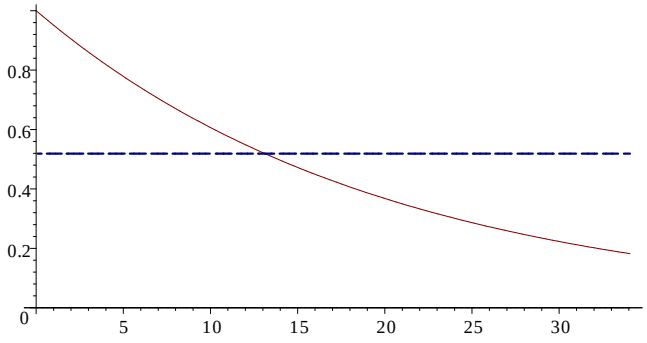

Figure 5: Real mortgage payments over time in a zero inflation environment (dashed line) and 5% inflation environment (curve).

Without inflation the real mortgage payments are constant, while in an inflationary environment the real mortgage payments decrease over time. In order to keep the real net present value the same in the two environments, the initial payments have to be higher in a world with non-zero inflation. That is, the real repayment profile is *tilted* towards the earlier periods. In other words, when inflation is high, the financial burden is "front-loaded" and the mortgage-payment-to-income-ratio is higher in the early years of the mortgage. Hence, liquidity constraints are more likely to bind and agents are less able to leverage. In turn, a more binding constraint in the first period of the mortgage depresses housing demand and prices. Note that if liquidity constraint were to be binding for a large set of agents, we would expect the price-rent ratio, and the mispricing measures, to be linked to movements in the real interest rate – we have seen in previous sections that this is not the case.

To test whether the tilt effect drives our results we perform two types of tests.

First, note that the intercept of the repayment scheme is proportional to the nominal interest rate, i, and that it is related to inflation only insofar as it affects the nominal interest rate. That is, once one controls for the nominal interest rate, inflation should not matter if the tilt effect is the driving force of our results. On the other hand, if the

mispricing is driven by money illusion, it should be related to inflation π as stressed in Equation (10).[20] Therefore, we regress our benchmark mispricing measures, ψ_t and ε, jointly on inflation and the nominal interest rate (both in levels and in logs). If the mispricing is driven by the tilt effect, inflation π should not play any role after controlling for the nominal interest rate i. This hypothesis is clearly rejected as shown in Table 3.

DepVar:	Regressors:			
	π_t	i_t	$\log(i_t)$	$\bar{R}^2$
$\hat{\psi}_t$	-3.13	-2.02		.85
	(6.395)	(2.096)		
$\hat{\psi}_t$	-3.30		$-.038$	.85
	(7.467)		(2.109)	
$\hat{\varepsilon}_t$	-3.27	-1.27		.65
	(3.498)	(.6548)		
$\hat{\varepsilon}_t$	-3.15		$-.034$	.66
	(3.616)		(.9055)	

Table 3: Regression coefficients and Newey and West (1987) corrected t-statistics in brackets.

The first row of Table 3 reports the multivariate regression of the ψ-mispricing on inflation and the nominal interest rate. The inclusion of the interest rate does not drive out the statistical significance of π and it increases the measure of fit by a mere 2% (see Table 1). Note also that inflation is highly statistically significant, while the interest rate is only significant at the 5% level. The second row uses $\log(i_t)$ instead of i_t and delivers almost identical results. The third and the fourth row use the ε-mispricing. In both cases inflation is highly statistically significant after controlling for the interest rate. Furthermore, both i_t and $\log(i_t)$ are not statistically significant and the inclusion of these regressors increase the measures of fit by no more than 1%.

These results reject the null that our key findings are simply capturing the tilt effect. Nevertheless, they could be due to a subset of the observations in the sample. To check for this, we perform the same exercise as reported in Table 3 using a rolling window with size equal to a third of the full sample.

Panel A and Panel B of Figure 6 depict the estimated regression coefficients and 95% confidence intervals of multi-variate regression of the ψ-mispricing on inflation π_t and the log of the nominal interest rate, $\log(i_t)$. The second row of Figure 6 reports the same analysis for the ε-mispricing. The point estimates are reported at the date of the last observation of the rolling sample (of 52 quarterly observations) – e.g., the point estimates corresponding to 1982:Q1 uses data from 1969:Q2 to 1982:Q1. Inflation is always significant, while $\log(i_t)$ is generally not statistically significant for the ψ-mispricing and never significant for the ε-mispricing. Moreover, the weak statistical

[20] If agents suffer from money illusion, the price-rent ratio comoves with the nominal interest rate, i (see Section 3.1) while, as pointed out in Section 3.2, the misspricing comoves with inflation, π.

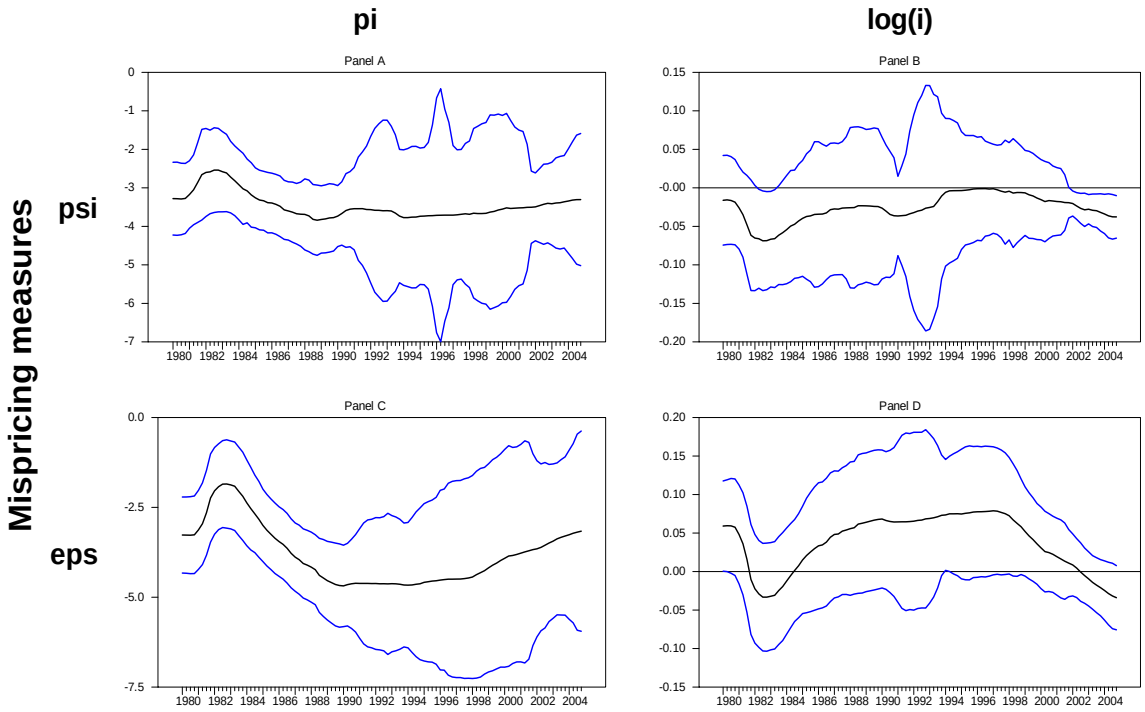

Figure 6: Multivariate regression of mispricing measures on inflation, π, and log interest rate, $\log(i)$, over rolling samples.

significance of $\log(i_t)$ in explaining the ψ-mispricing over the whole sample (Table 3) appears to be driven by the last four years of data. This last point is confirmed by an expanding window regression exercise (not reported here). Furthermore, the same qualitative results were obtained using the level of the nominal interest rate, i_t instead of $\log(i_t)$.

Second, we now present an alternative test to discriminate between the money illusion and the tilt-effect hypotheses based on the evolution of the mortgage market over time. Note that in our example the tilt effect arises since the nominal mortgage payments are constant, but more flexible mortgage contracts might reduce or eliminate it. Indeed, in the real world, agents can use multiple alternative financing schemes available on the market that are not affected by the tilt effect. For example this is the case for flexible interest rate mortgages, price level adjusted mortgages (PLAM) or the graduate payment mortgages (GPM).[21] This is especially true in the United Kingdom,

[21] On the other hand, Spiegel (2001) provides a rationale for endogenous credit rationing in the

where PLAM and GPM were available at least since the early 1970's. Furthermore, new, more flexible, mortgage products were introduced over the years in all major countries. In the U.S., for example, interest-only mortgages, which substantially lower the initial payments, have become very popular in recent years.[22] Hence, we would expect that the importance of the tilt effect – if it is there – declines over time. That is, the negative elasticity of the mispricing to inflation should become less negative over the sample period.

We empirically assess this hypothesis. Figure 7 depicts point estimates and Newey and West (1987) 95% confidence intervals of the univariate regressions of the estimated mispricing on π_t, i_t, and $1/i_t$ over a time-varying sample. We use the first ten years of data to obtain an initial estimate of the slope coefficient associated with each regressor, and we then add one data point at a time and update our estimates. For example, the point corresponding to 1992 first quarter is the estimated slope coefficient over the sample 1966 second quarter to 1992 first quarter.

Figure 7 Panel A, for the ψ-mispricing, and Panel D, for the ε-mispricing, reveal that the trend goes, if anything, in the opposite direction of what we would expect if the tilt effect were the driving mechanism behind the empirical link between housing prices and inflation. Over time, the negative relation between mispricing and inflation becomes more negative. The elasticity with respect to the interest rate is essentially flat (Panels B and E). Only the elasticity with respect to the log of the nominal interest rate reciprocal seems to decline at the end of the sample for the ψ-mispricing (Panel C), but this reduction is not statistically significant, while it is essentially flat for the ε-mispricing (Panel F). Overall, these findings suggest that it is unlikely that the tilt effect is the mechanism behind the empirical link between housing mispricing and inflation.

How do the findings in Figure 7 square with money illusion? Money illusion does not have a clear implication as to whether the elasticity of mispricing to inflation should vary over time. Nevertheless, the decline in the negative slope coefficient (in Panel A and D) is consistent with a setting in which households' attention to inflation depends on the recent history of inflation. Money illusion is very costly after and during a period of high inflation. Hence, households are more attentive to inflation and less prone to money illusion. In contrast, the cost of money illusion is perceived to be low after and during a period of low inflation – as in the last part of our sample – and hence money illusion is more wide-spread increasing the elasticity of the mispricing to inflation.

---

housing market due to moral hazard.

[22] See, for exampe, Lowenstein's article in the *The New York Times* on June 5, 2005, which cites the Lehman Brothers report "The Changing Landscape of the Mortgage Market" that describes the recent increase in interest-only mortgages.

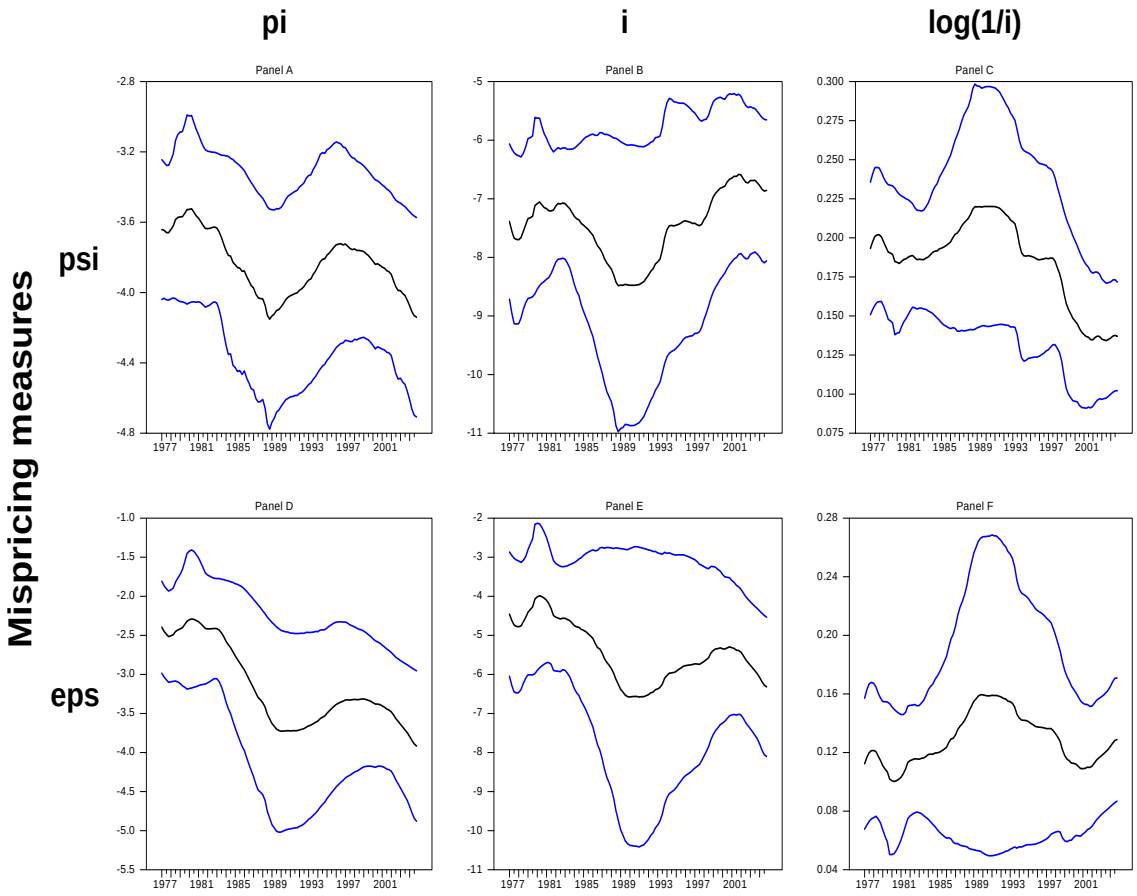

Figure 7: Point estimates and 95% Newey and West (1987) corrected confidence bounds of slope coefficients as sample size increases.

4.2 Lock-in Effect

When inflation and interest rates creep up, households that have secured a non-portable mortgage with a low fixed nominal interest rate in the past might be reluctant to buy a new, better quality house. This in turn could depress the demand for high-quality houses and lower the supply for low-quality houses. Therefore, given that the pools of rental houses and houses on the market for sale are not perfectly symmetric, this could be the driving force of the empirical link between the price-rent ratio and the nominal interest rate.

Noticing that this "lock-in effect" is asymmetric – since it predicts an additional reduction in housing demand for buying only when the interest rate is above the locked-in interest rate – we can perform a series of tests to assess this hypothesis.

First, we run a set of regressions of our mispricing measures, ψ and ε, on the interest rate and inflation interacted with a dummy variable that captures upward movements in the interest rate over the last four quarters.

That is, we run the following regressions:

$$\begin{aligned} \mu_t &= \hat{a}_i + \hat{b}_{i1} d_t i_t + \hat{b}_{i2} (1 - d_t) i_t + \hat{e}_{it} \\ \mu_t &= \hat{a}_\pi + \hat{b}_{\pi 1} d_t \pi_t + \hat{b}_{\pi 2} (1 - d_t) \pi_t + \hat{e}_{\pi t}, \end{aligned}$$

where μ_t is either $\hat{\psi}_t$ or $\hat{\varepsilon}_t$, and d_t is an indicator function of upward movements in the nominal interest rate i_t. The lock-in effect suggests that the coefficient estimate $\hat{b}_{\cdot 1}$ should be different from $\hat{b}_{\cdot 2}$. However, we cannot reject that $\hat{b}_{\cdot 1} = \hat{b}_{\cdot 2}$ for both regressions and both mispricing measures.

Second, using rolling samples (containing ten years of observations each) we test three separate hypotheses: $Corr\left[R_t^2, D_t\right] \neq 0$; $Corr\left[R_t^2, i_t\right] \neq 0$; and $Corr\left[R_t^2, \overline{p_t - l_t}\right] \neq 0$, where R_t^2 is the measure of fit of the regression of ψ on i in each rolling sample, D_t is the average d_t on a given subsample, and $\overline{p_t - l_t}$ is the average log price-rent ratio in a given subsample. All theses hypotheses are rejected at standard confidence levels.

These results may not be surprising since most mortgages are portable in the U.K.. This is unlike in the U.S. where a large share of the mortgage contracts are not portable. Nevertheless, as in the U.K., all stated hypotheses can be rejected with U.S. data also except for $Corr\left[R_t^2, D_t\right] \neq 0$.

5 Cross-regional Heterogeneity

In the previous section, we document money illusion as an aggregate phenomenon that can generate house price run ups without changes in economic fundamentals. In this section, we explore whether money illusion can be reconciled with the observed regional heterogeneity in price behavior within a country. In order to do this we shift our focus from the U.K. to the U.S. housing market because the cross-sectional heterogeneity is

more prominent in the U.S. and more regional data is available. We first document money illusion in the aggregate U.S. housing market. We then document heterogeneity in housing supply elasticity at the state level and explain how money illusion and heterogeneity in supply elasticity can rationalize the heterogeneity in regional price dynamics.

5.1 Aggregate U.S. Evidence

In this section, we examine the link between housing market mispricing measures and nominal values in the U.S. following the same procedure as in Section 3.2. The sample period available runs from 1970:Q1 to 2004:Q3. Univariate regression results are reported in Table 3. The first row shows that the proxies considered are all significant explanatory variables for the mispricing. Moreover, the sign of the estimated elasticity is the one we would expect under inflation illusion: the mispricing of the price-rent ratio tends to rise as inflation and nominal interest rates decrease. The coefficient estimates for the U.S. data are similar to the ones for the U.K. The measures of fit are somewhat smaller compared to the U.K. case, but this is likely to be due to the shorter sample period and poorer quality of U.S. data. An exception is the R^2 for the expected rent growth rate, which is higher for the U.S.

For a review of the measurement problems in U.S. data on housing, see McCarthy and Peach (2004). Nevertheless, the R^2 ranges from 28% when the explanatory variable is the nominal interest rate to 45% when we use inflation as the explanatory variable of the mispricing.

Dependent Variables:	Regressors:					
	π_t		i_t		$\log(1/i_t)$	
	coeff.	R^2	coeff.	R^2	coeff.	R^2
Panel A:						
$\hat{\psi}_t$	-6.65 (4.525)	.45	-6.30 (3.182)	.28	.141 (4.256)	.35
$\sum_{\tau=1}^{\infty} \rho^{\tau-1} \hat{E}_t \Delta l^e_{t+\tau}$	-2.87 (6.572)	.65	-3.46 (6.170)	.65	.066 (4.693)	.60
$-\sum_{\tau=1}^{\infty} \rho^{\tau-1} \tilde{E}_t r^e_{h,t+\tau}$	.76 (.211)	.01	4.65 (1.130)	.05	$-.066$ (.734)	.03
Panel B:						
$\hat{\varepsilon}_t$	-10.2 (5.148)	.48	-6.86 (2.648)	.15	.159 (3.238)	.21

Table 4: Univariate Regressions on inflation, nominal interest rate, and illusion proxy. Newey and West (1987) corrected t-statistics in brackets.

The second row of Table 4 shows that there is a significantly negative (positive) correlation between inflation and nominal interest rate (log of the nominal interest rate

reciprocal) and expected future rent excess growth rates. This could either be a consequence of a negative effect of inflation on the real economy or due to a higher degree of stickiness in housing rents than in the general price level. The regressors considered are able to explain between 60% to 65% of the time series variation in expected future growth rates. The last row shows that there is a no statistically significant link between inflation/nominal interest rate and future risk premia on housing investment. The coefficients for the U.S. are slightly lower compared to the U.K. coefficient, which is consistent with the different tax treatment of mortgage interest payments in both countries. Overall, these results imply a negative elasticity of the price-rent ratio to inflation (nominal interest rates) of about 8.7 (5.1) and that the largest contribution to this comes from the effect of inflation (nominal interest rate) on mispricing.

Table 5 reports the results of a Monte Carlo exercise (described in Section A.2 of the Appendix) analogous to the one presented in Section 3.2, which, as in the case of U.K. data, confirms the soundness of the empirical link between mispricing in the housing market and inflation, nominal interest rate, and the log of the nominal interest rate reciprocal. On the other hand, it shows that there is substantial uncertainty about the rational links between inflation (nominal interest rate) and the price-rent ratio, even though both variables show a significantly negative correlation with the risk premium on the housing investment.

DepVar:	Regressors:					
	π_t		i_t		$\log(1/i_t)$	
	coeff.	R^2	coeff.	R^2	coeff.	R^2
Panel A:						
$\hat{\psi}_t$	-6.06 [-7.32, -2.76]	.44 [.06, .66]	-5.84 [-7.12, -2.14]	.27 [.03, .66]	.130 [.070, .155]	.35 [.06, .60]
$\sum_{\tau=1}^{\infty} \rho^{\tau-1} \hat{E}_t \Delta l^e_{t+\tau}$	-2.86 [-8.17, 1.53]	.59 [.01, .96]	-3.45 [-7.27, -0.53]	.52 [.02, .71]	.066 [.003, .149]	.51 [.01, .70]
$-\sum_{\tau=1}^{\infty} \rho_t^{\tau-1} \tilde{E} r^e_{h,t+\tau}$	.44 [-4.84, 3.21]	.01 [0, .09]	4.23 [1.12, 5.82]	.04 [.01, .12]	$-.023$ [$-.097$, 0]	.07 [0, .15]
Panel B:						
$\hat{\varepsilon}_t$	-10.2 [-16.2, -7.25]	.48 [.36, .62]	-6.83 [-10, -4.79]	.15 [.11, .21]	.159 [.115, .25]	.21 [.16, .26]

Table 5: Median and 95% confidence intervals for slope coefficients and R^2. U.S. data.

5.2 The U.S. Regional Housing Markets

In the previous section we documented the presence of money illusion in the aggregate U.S. housing market. Yet house price shifts often vary significantly across different regions of the same country. In the U.S., the recent price increase seems to be much less pronounced in the Midwest compared to the coastal regions. In this section we

investigate whether money illusion can be reconciled with these heterogeneous price dynamics.

First, it should be mentioned that this regional heterogeneity is less extreme than it appears at first sight, once one separates property prices into land value and building value. Davis and Palumbo (2006) analyze 46 large metropolitan areas in the U.S. and find that the appreciation of residential land since the mid-1980s is a widespread phenomenon because the price of residential land has risen much faster than housing construction costs. The different rates of increase in land and building value could partially explain the observed cross-sectional heterogeneity in property price movements. To be more explicit, consider the following stylized example. Suppose in the low population density Midwest, 80% of the property value reflects the value of the building, while in New York City the building value is say only 10% of the property value. Then the same percentage increase in land price would cause a much larger increase in property value in New York City than in the Midwest.

Second, it is often argued that regional heterogeneity of property values is due to different housing supply elasticities. This could be another reason why money illusion impacts house prices across various regions differently. To understand how supply elasticity interacts with money illusion, let us consider a simple setting in which housing demand, $D(P, \pi)$, is decreasing in housing price, $D_1 := \partial D/\partial P < 0$, and inflation, $D_2 := \partial D/\partial \pi < 0$, due to money illusion. The supply of housing $S(P, X)$ is increasing in house prices, $S_1 := \partial S/\partial P > 0$, and also affected by cost shifters X. Applying the implicit function theorem on the market clearing condition to obtain $dP/d\pi$ and taking the derivative with respect to the slope of the supply function S_1, one obtains:

$$\frac{d(dP/d\pi)}{dS_1} = -\frac{D_2}{(D_1 - S_1)^2} > 0. \tag{16}$$

Since S_1 is inversely proportional to the elasticity of supply, this last equation shows that money illusion can generate large price movements in an area characterized by a low elasticity of supply (high S_1).[23] Therefore, we study whether S_1 changes across regional markets.[24] To do this, we nest our work in the housing supply literature and try to estimate the elasticity of new housing starts to house price changes. The existing literature has produced a wide range of both point estimates and conflicting methodologies [see DiPasquale (1999) for a survey]. Previous literature has focused on the empirical link between new housing starts and the housing price *level*. On the other hand, Mayer and Somerville (1996) argue that housing starts are a flow variable and therefore should be a function of other flow variables, and focus therefore on the link between housing starts and housing price *changes*. Also, using a panel of U.S.

[23] If builders also suffer from money illusion, this result still holds as long as households suffer from money illusion more than builders.

[24] In Glaeser, Gyourko, and Saks (2005), heterogeneity in supply elasticity is the endogenous outcome of households voting on zoning regulations.

quarterly state-level data over the period 1975:Q1–2005:Q4, we find that the OFHEO housing price indices seem to contain a unit root [as in Gallin (2003, 2004)], while this hypothesis can be rejected for the new housing starts series. As a consequence, we regress housing starts in state i, I_{it}, on the one-year price changes ΔP_{it}.

Note that our main interest is not to estimate the supply elasticity itself but to assess how it varies across different regions. In particular, we explore whether the relative degree of land availability generates heterogeneity in supply elasticity. For this purpose, we focus on an interaction term between prices, ΔP_{it}, and population density, pd_{it}. The previous literature has shown population density to be relevant in explaining cross-sectional housing price differences since it captures the relative scarcity of land [e.g., Voith (1996)].

Regressing new housing starts on price change, we would expect a positive coefficient for an upward-sloping supply curve. But, if high population density reduces the elasticity of supply, adding as a regressor price changes interacted with population density, $\Delta P_{it} \times pd_{it}$, we would expect a negative coefficient.

In order to distinguish movements along the supply curve from movements of the supply curve, we introduce as controls a set of construction-cost shifters. The additional regressors are the one-year change in the real ENR building cost index, the real interest rate on the three-month T-bill as a proxy for the cost of capital, and the one-year change in the per capita state-specific real wage as a proxy for labor cost changes.

In our panel regressions we also control for state fixed effects and state-specific cyclicality using state-specific quarterly dummies. In addition, we add to our regressions a third-order polynomial in a linear time trend and the time series of state population densities. We also add lagged new housing starts, I_{it-1}, to capture the high degree of persistence of the regressand. Population density is constructed removing water area from the total state area and the series is normalized to have unit mean. Nominal values are made real using the CPI less shelter. Unfortunately, we do not have data on land prices at the state level to add as an additional control. Following the previous literature, we include inflation to the regressions, and we also interact inflation with population density.

Since we have quarterly data over the period 1975:Q1–2005:Q4 for 50 states plus the District of Columbia, our sample size of $5,865$ observations implies that our regressions have more than $5,640$ degrees of freedom. Since prices, construction costs, and wages are potentially endogenously determined, we need to perform IV estimates. As instruments we use all the exogenous variables and their first lag, all the dummies and trends, and the first two lags of the endogenous variables. Weak instruments do not seem to be an issue since the measures of fit of the first-stage regressions range between 70% for price changes to 82% for the ENR building cost index. We also employ a heteroskedasticity and autocorrelation consistent estimator of the covariance matrix with Newey and West (1987) windows of 12 lags, and corrected to take into account the use of instrumental variables.

	Regressors:						$\bar{R}^2$
	I_{it-1}	ΔP_{it}	$\Delta P_{it} \times pd_{it}$	π_t	$\pi_t \times pd_{it}$	r_t	
(1)	.901	.843	−.064	−10.9		.172	.97
	(47.24)	(1.871)	(2.796)	(4.424)		(.0311)	
(2)	.901	.710		−11.26	.274	.195	.97
	(47.54)	(1.659)		(4.471)	(3.152)	(.0354)	
(3)	.901	.833	−.064	−11.14	.228	.140	.97
	(47.17)	(1.844)	(2.787)	(2.515)	(2.758)	(.0253)	

Table 6: Regressions of housing starts on its lag, yearly price-change, inflation and on interaction terms. Coefficients of additional controls are not reported. Corrected t-statistics in brackets.

Table 6 reports the point estimates of the main coefficients of interest in our panel regression. The first row shows, as expected, a positive correlation between price changes and new housing starts but the regression coefficient is significant only at the 10% level. Most importantly, the interaction term between price changes and population density is negative and highly statistically significant, suggesting that the elasticity of housing supply is reduced in areas where land is relatively scarce. Inflation shows a significantly negative relationship with housing starts as already found in Topel and Rosen (1988). This last finding is consistent with the proxy effect of inflation but also with a setting in which rational builders are aware that households suffer from money illusion. Indeed, Topel and Rosen (1988) find the effect of inflation on housing starts to be too strong to be explained by their rational investment model. The real interest rate is not significant. The second row replaces the interaction term $\Delta P_{it} \times pd_{it}$ with the interaction of inflation and population density, $\pi_t \times pd_{it}$. The point estimate of the other coefficients is almost unchanged and price changes are statistically significant only at the 10% level, while inflation is highly significant and the real interest rate is not. Interestingly, $\pi_t \times pd_{it}$ has a positive sign suggesting that inflation tends to reduce supply less in high density areas. The last row uses both interaction terms, and confirms the results of the previous regressions and also shows that the two interaction terms have different information content since they are both strongly significant even when employed jointly.

For robustness, we also performed regressions in which we modeled parametrically autoregression in the residuals, without obtaining significantly different results. Finally, we performed structural estimations of the investment model of Topel and Rosen (1988), but adding an interaction term between prices, inflation, and population density. Even though these last estimations regress housing starts on price levels, instead of price changes, we obtained qualitatively similar results.

Overall, the results in this section reconcile the presence of money illusion as an aggregate phenomenon and regional differences in price behavior.

6 Conclusion

This paper studies the link between inflation and housing prices. Our first-cut approach shows that the housing price-rent ratio is not affected by the real interest rate, but by the nominal interest rate. Moreover, we decompose time-series movements of the price-rent ratio into movements in a rational component and an implied mispricing component. We find that movements in inflation explain a large share of the time variation of the mispricing. These results are robust and hold for both the U.K. and the U.S. housing markets.

Two potential explanations of the link between the price-rent ratio and inflation naturally arise. First, inflation might make the economy riskier, or agents more risk averse, therefore increasing risk premia and driving down real estate prices. Second, current high inflation might be disruptive for the economy and/or inflation might proxy for future downturns, therefore depressing current housing value. We do not find supportive evidence for the first hypothesis, while the evidence in favor of the second hypothesis does not seem to be robust in the housing market context.

We also investigate possible explanations due to market frictions. First, inflation may tilt real mortgage payments towards the earlier years, making funding constraints potentially more binding. Second, an increase in inflation may dampen the demand for housing upgrades from individuals that have locked-in low nominal interest rates on an existing mortgages. Our extensive series of tests suggests that these market frictions are unlikely to be the mechanism behind the link between inflation and housing market mispricing. We also document substantial heterogeneity in housing supply elasticity across U.S. states due to differences in relative land scarcity, and we argue that this could reconcile our evidence of an aggregate money illusion phenomenon in housing markets with the observed heterogeneity in regional price behavior.

We therefore interpret our findings as supportive evidence for the money illusion hypothesis. Our findings provide a new argument in favor of price stability, since residential housing is the single largest asset class of households.

Several potential future research avenues come to mind. First, the analysis could be extended to a cross-section of countries to assess the role of institutional features. Preliminary results for Australia are also supportive of the money illusion hypothesis. Second, it would be interesting to study the common root of money illusion in markets as different as the residential housing market, the commercial real estate market and the stock and bond markets.

References

Abreu, D., and M. K. Brunnermeier, 2003, "Bubbles and Crashes," *Econometrica*, 71(1), 173–204.

Amihud, Y., 1996, "Unexpected Inflation and Stock Returns Revisited - Evidence from Israel," *Journal of Money, Credit, and Banking*, 28(1), 22–33.

Asness, C. S., 2000, "Stocks versus Bonds: Explaining the Equity Risk Premium," *Financial Analysts Journal*, 56(2), 96–113.

——— , 2003, "Fight the Fed Model," *Journal of Portfolio Management*, 30(1), 11–24.

Ayuso, J., and F. Restoy, 2003, "House Prices and Rents: An Equilibrium Asset Pricing Approach," Working paper, Bank of Spain.

Basak, S., and H. Yan, 2005, "Equilibrium Asset Prices and Investor Behavior in the Presence of Money Illusion," Working paper, London Business School.

Bauwens, L., M. Lubrano, and J.-F. Richard, 1999, *Bayesian Inference in Dynamic Econometric Models*. Oxofrd University Press, Oxford.

Bernanke, B., and M. Gertler, 1989, "Agency Costs, Net Worth, and Business Fluctuations," *American Economic Review*, 79(1), 14–31.

Bernanke, B., M. Gertler, and S. Gilchrist, 1996, "The Financial Accelerator and the Flight to Quality," *Review of Economics and Statistics*, 48, 1–15.

Blinder, A., 2000, "Comment on 'Near-Rational Wage and Price Setting and the Long-Run Phillips Curve'," *Brookings Papers on Economic Activity*, 2000(1), 50–55.

Boudoukh, J., and M. Richardson, 1993, "Stock Returns and Inflation: A Long-Horizon Perspective," *American Economic Review*, 83(5), 1346–1355.

Brunnermeier, M. K., 2007, "Comment on: Inflation Illusion, Credit and Asset Prices," in *Asset Prices and Monetary Policy*, ed. by J. Campbell. NBER, forthcoming.

Campbell, J. Y., 1991, "A Variance Decomposition for Stock Returns," *Economic Journal*, 101, 157–79.

Campbell, J. Y., and R. J. Shiller, 1988, "The Dividend-Price Ratio and Expectations of Future Dividends and Discount Factors," *Review of Financial Studies*, 1(3), 195–228.

Campbell, J. Y., and T. Vuolteenaho, 2004, "Inflation Illusion and Stock Prices," *American Economic Review Papers and Proceedings*, 94, 19–23.

Case, K. E., and R. J. Shiller, 1989, "The Efficiency of the Market for Single-Family Homes," *American Economic Review*, 79, 125–137.

——— , 1990, "Forecasting Prices and Excess Returns in the Housing Market," *AREUEA Journal*, 18, 253–273.

Clayton, J., 1996, "Rational Expectations, Market Fundamentals and Housing Price Volatility," *Real Estate Economics*, 24(4), 441–470.

Cohen, R. B., C. Polk, and T. Vuolteenaho, 2005, "Money Illusion in the Stock Market: The Modigliani-Cohn Hypothesis," *Quarterly Journal of Economics*, 120(2), 639–668.

Davis, M. A., and M. G. Palumbo, 2006, "The Price of Residential Land in Large U.S. Cities," mimeo, Federal Reserve Board.

DeLong, J. B., A. Shleifer, L. H. Summers, and R. J. Waldmann, 1990, "Noise Trader Risk in Financial Markets," *Journal of Political Economy*, 98(4), 703–738.

DiPasquale, D., 1999, "Why Don't We Know More about Housing Supply?," *Journal of Real Estate Finance and Economics*, 18, 9–23.

Earley, F., 1996, "Leap-Frogging in the UK Housing Market," *Housing Finance*, 32, 7–15.

Fama, E. F., 1981, "Stock Returns, Real Activity, Inflation and Money," *American Economic Review*, 71(4), 545–565.

Fama, E. F., and G. W. Schwert, 1977, "Asset Returns and Inflation," *Journal of Financial Economics*, 5(2), 115–146.

Fehr, E., and J.-R. Tyran, 2001, "Does Money Illusion Matter?," *American Economic Review*, 91(5), 1239–1262.

Ferson, W. E., S. Sarkissian, and T. Simin, 2002, "Spurious Regressions in Financial Economics?," NBER Working Paper 9143.

Fisher, I., 1928, *Money Illusion*. Adelphi, New York.

——— , 1930, *The Theory of Interest*. Maxmillan, New York.

Follain, J., 1982, "Does Inflation Affect Real Behaviour? The Case of Housing," *Southern Economic Journal*, 48, 570–582.

Froot, K. A., and M. Obstfeld, 1991, "Intrinsic Bubbles: The Case of Stock Prices," *American Economic Review*, 81(5), 1189–1214.

Gallin, J., 2003, "The Long-Run Relationship between House Prices and Income: Evidence from Local Housing Markets," Board of Governors of the Federal Reserve Discussion Paper 2003-17.

——— , 2004, "The Long-Run Relationship between House Prices and Rents," Board of Governors of the Federal Reserve Discussion Paper 2004-50.

Genesove, D., and C. Mayer, 2001, "Loss Aversion and Seller Behavior: Evidence from the Housing Market," *Quarterly Journal of Economics*, 116(4), 1233–1260.

Glaeser, E. L., J. Gyourko, and R. E. Saks, 2005, "Why Have Housing Prices Gone Up?," *American Economic Review Papers and Proceedings*, 95(2), 329–333.

Gultekin, N. B., 1983, "Stock Market Returns and Inflation: Evidence from Other Countries," *Journal of Finance*, 38(1), 49–65.

Gyourko, J., C. Mayer, and T. Sinai, 2006, "Superstar Cities," NBER Working Paper 12355.

Harrison, J. M., and D. Kreps, 1978, "Speculative Investor Behavior in a Stock Market with Heterogeneous Expectations," *Quarterly Journal of Economics*, 89, 323–336.

Hendry, D. F., 1984, "Econometric Modelling of House Prices in the United Kingdom," in *Econometrics and Quantitative Economics*, ed. by D. F. Hendry, and K. F. Wallis. Basil Balckwell, Oxford.

Himmelberg, C., C. Mayer, and T. Sinai, 2005, "Assessing High House Prices: Bubbles, Fundamentals, and Misperceptions," *Journal of Economic Perspectives*, 19(4), 67–92.

Hodrick, R. J., and E. C. Prescott, 1997, "Postwar U.S. Business Cycles: An Empirical Investigation," *Journal of Money, Credit, and Banking*, 29(1), 1–16.

Kearl, J., 1979, "Inflation, Mortgage, and Housing," *Journal of Political Economy*, 85(5), 1115–1138.

Kim, J.-Y., 1994, "Bayesian Asymptotic Theory in a Times Series Model with a Possible Nonstationary Process," *Econometric Theory*, 10(3), 764–773.

Kiyotaki, N., and J. Moore, 1997, "Credit Cycles," *Journal of Political Economy*, 105(2), 211–248.

Leontief, W., 1936, "The Fundamental Assumptions of Mr. Keynes' Monetary Theory of Unemployment," *Quarterly Journal of Economics*, 5(4), 192–197.

Lessard, D., and F. Modigliani, 1975, "Inflation and the Housing Market: Problems and Potential Solutions," *Sloan Management Review*, pp. 19–35.

Lintner, J., 1975, "Inflation and Security Returns," *Journal of Finance*, 30(2), 259–280.

Mankiw, G., and D. Weil, 1989, "The Baby Boom, the Baby Bust, and the Housing Market," *Regional Science and Urban Economics*, 19(2), 235–258.

Mayer, C. J., and C. T. Somerville, 1996, "Regional Housing Supply and Credit Constraints," *New England Economic Review*, pp. 39–51.

McCarthy, J., and R. W. Peach, 2004, "Are Home Prices the Next 'Bubble'?," *FRBNY Economic Policy Review*, 10(3), 1–17.

Modigliani, F., and R. Cohn, 1979, "Inflation, Rational Valuation and the Market," *Financial Analysts Journal*, 37(3), 24–44.

Muellbauer, J., and A. Murphy, 1997, "Booms and Busts in the UK Housing Market," *The Economic Journal*, 107, 1701–1727.

Newey, W. K., and K. D. West, 1987, "A Simple, Positive Semidefinite, Heteroskedasticity and Autocorrelation Consistent Covariance Matrix," *Econometrica*, 55, 703–708.

Ortalo-Magné, F., and S. Rady, 2006, "Housing Market Dynamics: On the Contribution of Income Shocks and Credit Constraints," *Review of Economic Studies*, 73(2), 459–485.

Patinkin, D., 1965, *Money, Interest, and Prices*. Harper and Row, New York.

Piazzesi, M., and M. Schneider, 2007, "Inflation Illusion, Credit and Asset Prices," in *Asset Prices and Monetary Policy*, ed. by J. Campbell. NBER, forthcoming.

Poterba, J. M., 1984, "Tax Subsidies to Owner-Occupied Housing: An Asset-Market Approach," *Quarterly Journal of Economics*, 99(4), 729–752.

——— , 1991, "House Price Dynamics: The Role of Tax Policy and Demography," *Brookings Papers on Economic Activity*, 1991(2), 143–203.

Ritter, J. R., and R. S. Warr, 2002, "The Decline of Inflation and the Bull Market of 1982-1999," *Journal of Financial and Quantitative Analysis*, 37(1), 29–61.

Scheinkman, J., and W. Xiong, 2003, "Overconfidence and Speculative Bubbles," *Journal of Political Economy*, 111(6), 1183–1219.

Schervish, M. J., 1995, *Theory of Statistics*. Springer, New York.

Shafir, E., P. Diamond, and A. Tversky, 1997, "Money Illusion," *Quarterly Journal of Economics*, 112(2), 341–374.

Sharpe, S. A., 2002, "Reexamining Stock Valuation and Inflation: The Implications of Analysts' Earning Forecasts," *Review of Economics and Statistics*, 84(4), 632–648.

Shiller, R. J., 1997a, "Public Resistance to Indexation: A Puzzle," *Brookings Papers on Economic Activity*, 1997(1), 159–228.

——— , 1997b, "Why Do People Dislike Inflation?," in *Reducing Inflation: Motivation and Strategy*, ed. by C. Romer, and D. H. Romer. University of Chicago Press, Chicago.

——— , 2005, *Irrational Exuberance*. Princeton University Press, Princeton and Oxford, 2nd edn.

Spiegel, M., 2001, "Housing Return and Construction Cycles," *Real Estate Economics*, 29(4), 521–551.

Stein, J. C., 1995, "Prices and Trading Volume in the Housing Market: A Model with Downpayment Effects," *Quarterly Journal of Economics*, 110(2), 379–406.

Thaler, R., 1980, "Toward a Positive Theory of Consumer Choice," *Journal of Economic Behavior and Organization*, 1.

Tobin, J., 1972, "Inflation and Unemployment," *American Economic Review*, 62(1), 1–18.

Topel, R., and S. Rosen, 1988, "Housing Investment in the United States," *Journal of Political Economy*, 96(4), 718–740.

Torous, W., R. Valkanov, and S. Yan, 2005, "On Predicting Stock Returns with Nearly Integrated Explanatory Variables," *Journal of Business*, 78(1), forthcoming.

Tsatsaronis, K., and H. Zhu, 2004, "What Drives Housing Price Dynamics: Corss-country Evidence," *BIS Quarterly Review*, pp. 65–78.

Tucker, D. P., 1975, "The Variable-Rate Graduated-Payment Mortgage," *Real Estate Review*, pp. 71–80.

Tversky, A., and D. Kahneman, 1981, "The Framing of Decisions and the Psychology of Choice," *Science*, 211, 453–458.

Voith, R., 1996, "The Suburban Housing Market: The Effects of City and Suburban Job Growth," *Federal Reserve Bank of Philadelphia Business Review*, pp. 1–13.

Wheaton, W. C., 1985, "Life-Cycle Theory, Inflation, and the Demand for Housing," *Journal of Urban Economics*, 18, 161–179.

Zellner, A., 1971, *An Introduction to Bayesian Inference in Econometrics*. Wiley, New York.

A Appendix

A.1 Data Description

A.1.1 U.K. Data

The housing price series is from the Nationwide Building Society, and covers the sample period 1966:Q2 to 2005:Q1. Over the period 1966:Q2 to 2005:Q5, the index is constructed as a weighted average using floor area, therefore allowing to control for the influence of house size. Over the periods 1974:Q1 to 1982:Q4, and 1983:Q1 to 1992:Q1 additional controls (for region, property type, etc.) have been added in the construction of the index. Since 1993, the index also takes into account changes in the neighborhood classification. The rent series is constructed combining several sources available through the Office of National Statistics. Over the period 1966:01 to 1987:01, we use the CTMK LA:HRA series of rents on dwellings paid by tenants in the U.K. and we combine it with the data on the stock of housing available through the Office of the Deputy Prime Minister. Over the period 1987:02 to 1987:12, we use the RPI-SGPE rent series of monthly percent changes over one month. Over the period 1988:01 to 2005:02, we use the CZCQ - RPI series of percentage changes in rent over one year. The rent-free tenancies are excluded from the calculation of average rents. To obtain a series in levels for the price-rent ratio, we scale the index series to match the level of the average price-rent ratio in 1990. As interest rate we use the twenty-year par yield on British Government Securities available over the sample 1963:Q4 to 2004:Q4. All the results presented in the paper are based on the longest possible sample given the data at hand (1966:Q2 to 2004:Q4). The cross-regional price variability is computed as the variance of the Nationwide quarterly log price index across the main fourteen macro regions of the U.K. available over the period 1973:Q3 to 2005:Q4. Total within country migration probability is computed over the period 1975:Q1 to 2004:Q4 from the NHSCR quarterly tables using total inflows and outflows of England, Wales, Scotland, and Northern Ireland normalized total U.K. resident population.

The implied inflation series, available over the period 1982:Q1 to 2005:Q1, is from the Bank of England and is constructed using the inflation-protected ten-years government securities.

The real GDP measure is the seasonally adjusted, chained volume measures with constant 2002 prices and is available over the period 1955:Q1 to 2005:Q1 from the Office of National Statistics.

A.1.2 U.S. Data

Aggregate Data. To construct the housing price index series, we use (i) the weighted repeat-sale housing price index from the Office of Federal Housing Enterprise Oversight over the sub-sample 1976:01 to 2004:03 and the (ii) Census Bureau housing price index

(obtained through the Bank of International Settlements) over the period 1970:01 to 1975:04. To construct the rent index, we use the CPI-Rent from the Bureau of Labor Statistics. We re-scale the indexes to levels to match the historical average of the U.S. price-rent ratio over the same sample [as reported in Ayuso and Restoy (2003)]. As long-run interest rate, we use the return on the ten-year Treasury bill. As measure of inflation, we use the CPI without housing.

Regional Data. As housing price index, we use the quarterly OFHEO Housing Price Index (HPI) for 50 states and the District of Columbia over the period 1975:Q1 to 2005:Q4. The availability of these data series determines the sample of the panel analysis. As new housing starts measure, we use the quarterly not seasonally adjusted state level Private Housing Units Permits Authorized series from the Markets database available through Global Insight. The labor cost measure is the state level quarterly total Wages and Salary series provided by the Bureau of Economic Analysis. As state population measure, we use the yearly total resident population estimates provided by the Bureau of Census of the U.S. Department of Commerce. The series are interpolated at quarterly frequency. We compute state land area as total area minus water area from the Census 2000 data. As building cost proxy we use the aggregate quarterly not seasonally-adjusted Engineering News Record (ENR) Building Cost Index (BCI) average of twenty U.S. cities. The BCI is a weighted index of skilled labor, structural steel shapes, portland cement, and lumber costs. Nominal values are made real using the CPI less shelter price index.

A.2 Assessing Uncertainty

To assess uncertainty in the regression results in Table 2, we report 95% confidence intervals for the estimated slope coefficients and R^2constructed via Monte Carlo integration by drawing form the posterior distribution of the estimated VAR coefficients. We proceed as follows:

1. We draw covariance matrices $\grave{\Sigma}$ from the inverse Wishart with parameters $\left(n\hat{\Sigma}\right)^{-1}$ and $n-m$.

2. Conditional on $\grave{\Sigma}$ we draw a vector of coefficients for the VAR, $\grave{\beta}$, from

$$\grave{\beta} \sim N\left(\hat{\beta}, \grave{\Sigma} \otimes (X'X)^{-1}\right).$$

3. Using the draws of the VAR slope coefficients, $\grave{\beta}$, we construct expected discounted sums of rent excess growth rates ($\sum_{\tau=1}^{\infty} \rho^{\tau-1} \grave{E}_t \Delta l^e_{t+\tau}$) and obtain the excess housing returns ($\sum_{\tau=1}^{\infty} \rho^{\tau-1} \grave{E}_t r^e_{h,t+\tau}$) in order to compute pricing errors $\grave{\psi}$ and $\grave{\varepsilon}$.

4. We then regress $\grave{\psi}_t$, $\grave{\varepsilon}$, $\sum_{\tau=1}^{\infty} \rho^{\tau-1} \grave{E}_t \Delta l^e_{t+\tau}$ and $\sum_{\tau=1}^{\infty} \rho_t^{\tau-1} \grave{E}_t r^e_{h,t+\tau}$ on π_t, i_t, and the log of the inflation-biased evaluation $1/i_t$, and we store the estimated slope coefficients and measures of fit.

5. We repeat this procedure 10,000 times and compute confidence intervals for the OLS slope coefficients associated with π_t, i_t and the log of $1/i_t$, and for the corresponding measures of fit, from the corresponding percentiles of the Monte Carlo iterations.

A.3 Assessing the Role of the Business Cycle

To construct a business cycle proxy for the U.K. we follow Hodrick and Prescott (1997), that is we estimated the following state-space model:

$$\begin{aligned} \Delta y_t &= g_t + c_t \\ g_t &= 2g_{t-1} - g_{t-2} + v_t \end{aligned} \tag{17}$$

where Δy_t is GDP growth from quarter $t-5$ to quarter t, g_t is the unobserved state variable meant to capture the smooth time varying trend, and c_t is the cyclical component. The variance of v_t is normalized to be $1/1,600$ times the variance of the cyclical component, c_t, as it is customary with quarterly data. This state-space representation is estimated via Kalman filter and Kalman smoother.

Figure 8 plots the time series of the U.K. exponentially smoothed quarterly inflation, the return on the twenty-year government bonds, and the Hodrick and Prescott (1997) filtered estimate of the GDP business cycle. The HP-estimate seems to capture fairly well the business cycle over the period considered. Moreover, there is no clear comovement between inflation and the business cycle.

Table A1 reports OLS regressions of our mispricing measures ($\hat{\varepsilon}_t$ and $\hat{\psi}_t$) on the variables meant to capture money illusion (π_t, i_t and $\log(1/i)$) and the business cycle component of GDP identified by the H-P filter ($\hat{c}_t$).

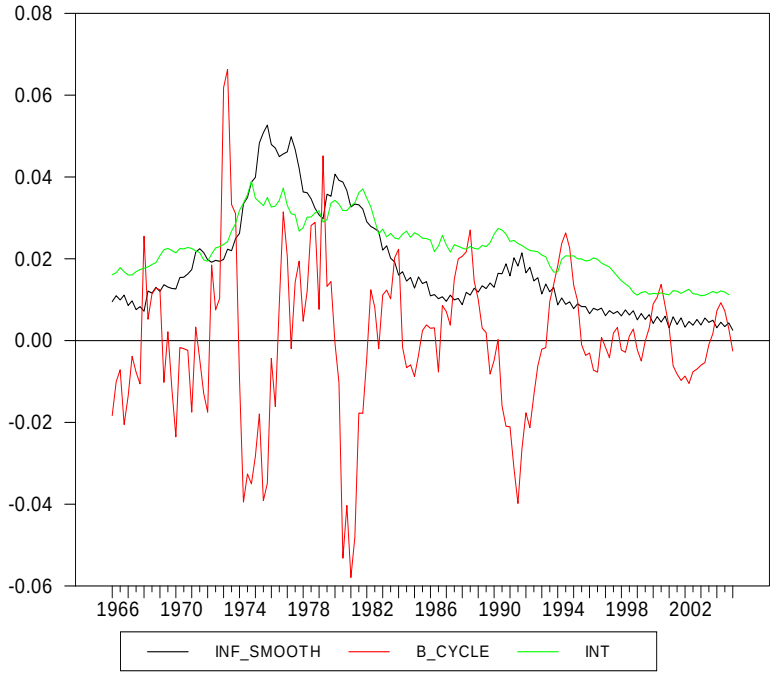

Figure 8: U.K. business cycle and inflation

Row:	Dep. variable:	Regressors: $\hat{c}_t$	π_t	i_t	$\log(1/i)$	R^2
(1)	$\hat{\psi}_t$	0.81 (1.959)				.07
(2)		0.32 (2.135)	−4.00 (13.761)			.85
(3)		0.378 (2.168)		−6.64 (11.137)		.76
(4)		0.50 (2.590)			0.13 (8.121)	.71
(5)	$\hat{\psi}'_t$	1.11 (0.963)				.01
(6)		0.36 (0.349)	−5.98 (2.279)			.17
(7)		0.41 (0.369)		−10.5 (2.436)		.17
(8)		0.55 (0.503)			0.24 (2.605)	.19
(9)	$\hat{\varepsilon}_t$	0.85 (2.201)				.07
(10)		0.41 (2.281)	−3.80 (7.801)			.67
(11)		0.49 (2.158)		−6.10 (6.399)		.57
(12)		0.60 (2.462)			0.12 (5.769)	.56

Table 7. Regressions on business cycle fluctuations, inflation, nominal interest rate, and illusion proxy. Newey and West (1987) corrected t-statistics in brackets.

It is clear from the first and fifth rows of Table 7 that the business cycle has little (in the case of $\hat{\varepsilon}_t$) or no (in the case of $\hat{\psi}_t$) explanatory power for the mispricing. The remaining rows clearly show that the inclusion of the business cycle in the OLS regressions for the mispricing a) does not drive out the statistical significance of π_t, i_t and $\log(1/i_t)$, b) does not significantly change the point estimates of the elasticities of the mispricing reported in Table 1, and c) does not increase significantly our ability to explain the time variation in the mispricing (comparing Table 7 to Table 1, we have that the increase in R^2 is ranges from 0% to 4% percent, and there is virtually no increases in the – non reported – $\bar{R}^2$).