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DRIVERS OF HOUSING CONSTRUCTION: A EUROPEAN COMPARISON

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ABSTRACT

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Drivers of housing construction: A European comparison*

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Abstract

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JEL Classifications: R31; E22; E32; E52; C22;.

Key Words: Housing supply; Building permits; ARDL modeling; Monetary policy transmission; Construction costs; European housing markets.

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1 Introduction

Residential construction is central to both macroeconomic stability and housing affordability in advanced economies. In the European Union, the broader construction sector contributes approximately to 5.6% of value-added and employs around 6.7% of the total workforce, with residential investment representing a substantial share of total investment (Eurostat, 2025, p. 51). Beyond its direct contribution to output and employment, residential construction amplifies business cycles and determines the supply of housing units: expanding construction exerts downward pressure on house prices and rents, while contracting supply intensifies scarcity. Understanding the drivers of construction activity is therefore paramount for policymakers seeking to stabilize the business cycle while ensuring adequate housing supply.

The recent monetary tightening cycle has brought these concerns into sharp focus. Following a decade of accommodative policy during which mortgage rates fell below 1%, the European Central Bank began raising rates rapidly in July 2022, reaching approximately 3.5–4.0% by late 2023—the fastest tightening in ECB history. The effects on residential construction have been severe and heterogeneous. In Germany, building permits approximately halved between their 2021 peak and mid-2024[1], while construction prices rose by over 16% year-on-year in 2022. In France, permits declined by nearly 30%, and Belgium experienced a roughly 50% drop from mid-2021 to September 2023. These developments have raised concerns about a construction collapse that could exacerbate housing shortages and erode sector capacity for years to come (Jonas et al., 2023; Leiss and Wohlrabe, 2024). At the European level, the European Systemic Risk Board issued warnings to several member states in 2022, identifying house price overvaluation as a key vulnerability (Ciocchetta et al., 2024)—a concern now compounded by declining construction activity.

Despite policy urgency, the empirical literature on construction supply determinant faces significant challenges. First, as DiPasquale (1999) emphasizes, aggregate studies consistently encounter puzzling empirical patterns, most notably, the persistent failure of construction cost coefficients to confirm the theoretically expected negative sign. Second, existing studies focus on single countries or employ panel methods that impose homogeneous coefficients across countries, potentially obscuring important heterogeneity in supply responses (Duca et

[1] Building permits precede actual construction activity by approximately 12–24 months (Leiss and Wohlrabe, 2024).

al., 2021). Third, recent literature demonstrates that housing supply cannot be understood in isolation from financial conditions and rental market dynamics (Paz-Pardo et al., 2024), suggesting that a comprehensive framework must account for multiple transmission channels simultaneously. This paper addresses these gaps by providing a systematic multi-country analysis of residential building permit determinants for five major European economies, namely Germany, France, the Netherlands, Belgium, and Austria, using monthly data from January 2000 through June 2024, with selected variables interpolated from quarterly sources.

Building permits capture early-stage developer decisions that are potentially more sensitive to financing conditions and forward-looking expectations than residential investment or completions. Our sample encompasses the unprecedented 2022–2024 tightening cycle, providing crucial identifying variation for the financing cost channel. For each country, we estimate six nested ARDL error-correction model (ARDL-ECM) specifications with bootstrap inference, progressively adding demand-side, financing cost, supply-side, and asset market variables to assess coefficient stability and mediation patterns. This extends the primarily US-focused literature reviewed by Duca et al. (2021) and complements single-country studies such as Jonas et al. (2023) for Germany and Ciocchetta et al. (2024) for European house prices.

Our analysis yields three main findings. First, we document substantial cross-country heterogeneity that reveals two distinct regimes shaped by institutional differences in mortgage markets, financial depth, and rental market structure. Germany is subject to a "fundamental and supply-side driven" regime where construction responds to financing costs, construction costs, unemployment rate and, notably, rental income rather than house price momentum. France and the Netherlands display "speculative and demand-side driven" regimes dominated by income elasticities and house price acceleration, albeit likely through structurally different transmission channels: credit amplification in the Netherlands versus institutionally created investment incentives in France. Belgium occupies an intermediate position, while Austria fails to establish cointegration in any specification, a finding we attribute to its dominant non-profit housing sector and subsidized credit system that insulate construction from market-based transmission.

Second, we shed new light on the persistent construction cost puzzle identified by DiPasquale (1999). The counter-intuitive positive coefficients on construction cost inflation we find for France and the Netherlands may reflect anticipatory investment behavior: when developers extrapolate recent cost increases into the future, while prospective homeowners extrapolate house price increases, accelerating construction becomes rational.

Third, dynamic multipliers reveal a consistent hierarchy of adjustment speeds across determinants: income and house price effects materialize within 2–4 months, financing cost effects require 4–8 months, and unemployment and rent effects display the longest lags (8–13 months). These adjustment periods imply that permit responses operate at business-cycle frequencies, with direct implications for the timing of policy interventions.

Our findings carry direct policy relevance. The strong financing cost effects in Germany and France, combined with evidence of construction sector distress (Leiss and Wohlrabe, 2024), support arguments for counter-cyclical public investment to maintain sector capacity during tightening cycles. At the same time, the heterogeneous responses imply that uniform European housing policies are unlikely to be effective: subsidized financing may work in Germany but have limited impact in the Netherlands, where speculative price dynamics dominate the cost-of-capital channel.

The remainder of the paper proceeds as follows. Section 2 presents the theoretical framework. Section 3 describes the data and econometric methodology. Section 4 presents empirical results, including long-run coefficient estimates, cross-country comparisons, and dynamic multipliers. Section 5 discusses policy implications and concludes.

2 A comprehensive model of the housing market

Our empirical framework builds on the four-quadrant model developed by DiPasquale and Wheaton (1992) and DiPasquale and Wheaton (1994), which provides theoretical justification for modeling residential construction as a function of macroeconomic variables. The model links four interconnected components: the space market (determining rents), the asset market (determining property values), the construction sector (determining new supply), and the stock adjustment mechanism linking construction flows to the existing housing stock. Following the classification proposed by Melecky and Paksi (2024), construction activity determinants can be grouped into demand-side, financing cost, and supply-side variables. Following these references, Table 1 summarizes the predicted effects and their transmission channels: it serves as a reference point for the empirical analysis throughout the paper.

Variable	Symbol	Classification	Primary transmission channel	Sign
Income	y	Demand-side	Space-market → rents → asset market	+
Unemployment	u	Demand-side	Income security, credit access	−
Interest rate	i	Financing cost	Cost of capital, developer funding	−
Cost of construction	c	Supply-side	Marginal cost of new housing units	−
House prices	h	Asset market	Developer profitability signal	+
Rents	r	Rental market	Buy-to-let profitability signal Insufficient equity/regulation constraints	+ −

Notes: Predicted signs derived from the DiPasquale-Wheaton four-quadrant model (DiPasquale and Wheaton, 1992; DiPasquale and Wheaton, 1994). House prices and rents are endogenous market outcomes that transmit signals from multiple categories simultaneously. The ambiguous sign for rents reflects the theoretical tension between the buy-to-let investment channel (positive) and potential market tightness, insufficient equity or rent regulation constraints (negative), as discussed by Paz-Pardo et al. (2024). Classification follows Melecky and Paksi (2024).

Table 1: Theoretical Predictions For Building Permit Determinants

A crucial insight distinguishing this framework from simpler investment models is its self-correcting mechanism. Rising demand generates higher property prices and temporarily stimulates construction. However, as the housing stock expands, land prices—which depend on the stock rather than on flows—rise endogenously and absorb excess returns from above equilibrium house prices, returning construction to its equilibrium rate.

Interest rates (i), the key financing cost variable, transmit to construction through two channels: increasing the cost of capital in the asset valuation equation and raising developer funding costs. Income (y) shifts housing demand in the space market[2], raising both rents and property values through capitalization. House prices (h) serve as the primary profitability signal for developers: higher prices increase expected returns from construction projects relative to development costs. This relationship should be self-correcting as previously mentioned. Crucially, however, the mechanism operates primarily through commercial developers rather than self-building households, who build for own occupancy rather than for capital gains, and for whom house price appreciation represents a cost comparison with existing stock rather than a profitability signal.[3] This compositional difference generates a prior that house price elasticities of building permits should be stronger in developer-dominated markets—a prediction we revisit in the cross-country comparison in Section 4.

[2]This term refers to the usage of real property, in contrast to the real estate's role in the asset market.

[3]The relative weight of these two groups of economic agents varies across our sample countries: Germany has a comparatively large share of self-build construction, particularly in the single-family segment, while France and the Netherlands are more strongly dominated by commercial developers, especially in multi-family housing. To the best of our knowledge, Belgium occupies an intermediate position.

Construction costs (c), the key supply-side variable, affect construction directly by shifting the marginal cost of producing new housing units. Unemployment (u) proxies for labor market strength and broader economic conditions, affecting construction through reduced household income security and tighter credit access.

Rental prices (r) transmit to construction through two opposing channels, generating theoretical ambiguity. On the one hand, the positive channel operates via buy-to-let investment (Paz-Pardo et al., 2024): higher rents raise expected rental yields and stimulate developer activity, particularly in multi-family construction where rental income directly determines project returns. On the other hand, the negative channel reflects market tightness: rising rents may signal a capacity-constrained environment with low vacancy rates and elevated construction costs, in which new supply responds sluggishly regardless of the profitability signal, particularly when there are equity and liquidity constraints. Alternatively, rising rents may themselves be a consequence of households being priced out of homeownership by tight credit conditions. Here, rents are a counter-cyclical indicator of a weakening ownership market rather than a positive investment signal. The net effect depends on which transmission channel dominates.

The model also distinguishes between adjustment horizons (DiPasquale and Wheaton, 1994). In the short-run, the housing stock is fixed and adjustment occurs through price changes in the space and asset market. In the long run, the stock adjusts: reduced construction, for instance, lowers the housing stock, supporting higher rents and partially offsetting the initial price decline, illustrating the self-correcting mechanism. Overall, this stock-flow adjustment process, where construction represents a flow variable targeting a stock equilibrium, provides the theoretical foundation for the ARDL-ECM framework developed in Section 3.4.

3 Data and econometric analysis

First, we describe the details of the dataset, before explaining how to use it for our analysis of construction activity along the lines of the aforementioned model.

3.1 Data description

We use monthly data from January 2000 to June 2024 for five European countries: Germany, France, Netherlands, Belgium, and Austria. Our primary dependent variable is the number of building permits for residential buildings (excluding residences for communities), which captures early-stage construction decisions before capital is committed in residential investment. Lerbs (2014), studying German county-level housing supply, similarly uses construction permits as the dependent variable, noting that permits "represent supply intentions more accurately than housing completions" (p. 4). Table 2 provides detailed variable definitions and data sources.

Variable	Description of the original data	Price term	Source
p	Building permits, number of dwellings, residential buildings, except residences for communities, 2021=100, SA	—	Eurostat
h	House prices, const. prices, SA	real	OECD
c	Producer prices, new residential building, except residences for communities, deflated by HCPI, 2021=100, SA	real	Eurostat
y	Gross disposable income of households per capita, calendar adjusted, deflated by HCPI, EUR, SA	real	Eurostat
u	Unemployment rate of all persons and all ages as a percentage of active population, SA	—	Eurostat
r	Rent prices, deflated by HCPI, SA	real	OECD
i	Interest rates, MFI interest rates, households & NPISH, lending for house purchase excluding revolving loans & overdrafts, convenience & extended credit card debt [A22-A2Z], annualised agreed rate (AAR) / narrowly defined effective rate (NDER), all maturities, total, new business	nominal	ECB

Note: All index variables were re-based to 2015 = 100 for consistency across countries and sources. All variables are in real terms with the exception of interest rates. For consistency reason, using nominal rates may be criticized. On the other hand, many empirical studies treat nominal rates as relevant driver of real output responses, i.a. Sims (1992). A data-driven rationale stems from the observability in loan contracts.

Table 2: Data Description

All data are obtained from Eurostat, the OECD, and the European Central Bank. Real house prices (h), real rent prices (r) and real disposable household income per capita (y) are available only at quarterly frequency and are therefore converted to monthly using cubic interpolation. Building permits (p), real construction costs (c), unemployment rates (u), and nominal interest rates on new housing loans (i) are available at monthly frequency.

Although some series are available for longer periods in certain countries, we restrict the analysis to 2000m1–2024m6 for all countries to ensure comparability across specifications

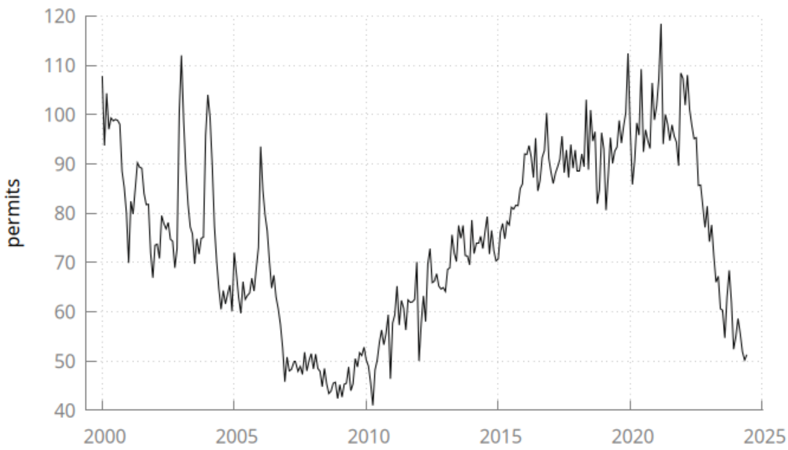

Figure 1: Time-Series of Permits in Germany Showing Excessive Temporary Values in 2003, 2004 and 2006.

within each country and consistency in cross-country comparisons. This period captures multiple phases of housing cycles, including the boom before the Global Financial Crisis (GFC), the GFC, the European sovereign debt crisis, and the recent interest rate tightening as a reaction to the energy price shocks and the inflationary episodes since 2021/22.

The original permits series for Germany exhibits implausibly high values in 2003, 2004, and 2006 (Figure 1), reflecting two distinct anticipatory effects. The values in 2003–2005 reflect the reduction (2004) and subsequent abolition (January 2006) of the "Eigenheimzulage", Germany's principal homeowner construction subsidy. The spike in 2006, by contrast, is more plausibly attributable to the value-added tax increase from 16% to 19% that took effect on 1st January 2007 and was announced in the 2005 coalition agreement. Because construction services are subject to the standard VAT rate, developers and households had strong incentives to file permit applications before the tax increase. Figure 1 illustrates the German case, but similar—though less pronounced—temporary spikes are observed in the remaining countries[4], likely driven by analogous country-specific policy changes or statistical re-classifications. We smooth these policy-driven spikes by replacing the affected months with the average of the three preceding and three subsequent months. This local-mean adjustment is designed to neutralize short-lived, pre-announced policy-induced pull-forward effects such as the VAT increase announced in late 2005 and effective in January 2007. This prevents transitory distortions from biasing the estimated long-run relationships.

[4]Analogous anticipation effects occurred around, for instance, successive reforms of the French Pinel/Duflot rental-investment tax credits and changes to Dutch LTV limits and transfer-tax rules, supporting the general case for spike smoothing prior to estimation.

Our focus on building permits rather than residential investment follows recent housing market literature emphasizing the importance of early-stage decisions (Cavalleri et al., 2019). Building permits offer three advantages for studying construction determinants: (i) they capture forward-looking expectations about future demand; (ii) they can adjust more rapidly than actual investment, providing higher-frequency information on construction sentiment; and (iii) they allow more precise identification of responses to macroeconomic shocks than quarterly investment data. The main disadvantage is that not all approved permits result in actual construction due to cancellations or postponements. However, the permit-to-completion relation is quite strong. The correlation between annual building permits and residential investment is $\rho = 0.78$ for Germany and above $\rho = 0.70$ for all other countries in our sample.

3.2 Descriptive statistics and visualization

Table 3 presents summary statistics for year-over-year changes in key variables across our five countries. Several patterns emerge that motivate our econometric analysis. Building permits exhibit negative mean growth in all countries except Belgium, with Germany showing the largest average decline (-2.2%). Volatility is substantial and heterogeneous, with the Netherlands displaying the highest standard deviation (31.3%), followed by Belgium (21.4%) and France (19.1%).

House prices show positive mean growth across all countries, ranging from 0.5% (Germany) to 2.5% (France), with Belgium (2.0%) and Austria (2.0%) exhibiting similar dynamics. Volatility patterns mirror those of permits, with the Netherlands (5.3%) showing the highest variability, followed by France (5.5%) and Austria (4.4%). Income growth is consistently positive but modest across countries, averaging between 0.4% (Austria) and 1.0% (France). In this case, standard deviations remain relatively small (1.0-2.0%).

Interest rates display universally negative mean changes, reflecting the downward trend in borrowing costs over our sample period. Germany shows the strongest average decline (-0.13 percentage points), while Austria exhibits the smallest (-0.08 percentage points). This secular decline in financing costs provides important variation for identifying interest rate effects on construction, particularly when combined with the sharp reversal during 2022–2024 included in our sample. Construction costs show positive mean growth ranging from

0.4% (Belgium) to 1.0% (Germany, Austria), with the Netherlands exhibiting exceptionally high volatility (5.4). Unemployment trends are generally negative, with Germany showing the strongest average decline (-0.20 percentage points). Austria stands out as the only country with slight positive mean growth (0.05), while the Netherlands shows near-zero change (-0.001). The heterogeneous unemployment dynamics suggest that labor market conditions affect construction through country-specific channels, see for instance Carrillo-Tudela et al. (2025).

– Place Table 3 here.

Figure 2 visualizes these patterns in levels, revealing substantial cross-country heterogeneity in both trends and cyclical fluctuations. Notably, the majority of series display marked volatility around the 2008–2009 financial crisis and again during 2020–2024 due to the COVID-19 pandemic and the recent interest rate tightening in the wake of energy price shocks following Russia's attack on Ukraine. This recent period deserves special attention and provides crucial identifying variation for our analysis, as it represents the first substantial increase in financing costs after more than a decade of declining rates.

– Place Figure 2 here.

Table 4 reports contemporaneous, unconditional correlations on year-over-year changes which do not imply causality nor capture lag structures. They nonetheless motivate our ARDL-ECM approach, which allows for dynamic and long-run relationships. Again, these patterns suggest a high degree of cross-country heterogeneity, which will be analyzed in the following sections. A summary of the most notable correlations can be found in the Appendix.

– Place Table 4 here.

3.3 Unit root properties of the core variables

We test for unit roots to verify the integration order of our variables, as the ARDL bounds-testing framework permits a mix of $I(0)$ and $I(1)$ regressors but not $I(2)$ (Pesaran et al., 2001). For each country and variable, we apply the ADF-GLS test (Elliott et al., 1996) and the KPSS test (Kwiatkowski et al., 1992). The ADF-GLS tests the null hypothesis of a unit root, while the KPSS tests the null of stationarity (level- or trend-stationarity). Results are reported in Tables 5 and 6.

– Place Table 5 here. –

– Place Table 6 here. –

For the logarithm of building permits ($\ln p$), the ADF-GLS test fails to reject the unit root null in levels for most countries (p-values ranging from 0.133 to 0.969) with the only exception being Belgium (p-value 0.027). First differences show strong evidence of stationarity for France, Netherlands, Belgium, and Austria (p-values ≤ 0.028). Germany shows weaker evidence in first differences (p-value 0.580), likely reflecting some kind of mean shifts between 2004 and 2008 as documented in Figure 1. The KPSS test uniformly rejects stationarity in levels but fails to reject in first differences across all countries. We therefore treat building permits as $I(1)$.

Logarithm of real income ($\ln y$), unemployment (u), and the interest rate (i) display similar patterns. All three variables show high ADF-GLS p-values in levels (mostly above 0.40), indicating unit roots, while first differences generally show evidence of stationarity. The KPSS test confirms this pattern: all three variables are non-stationary in levels but stationary in first differences for most countries. We therefore treat income, unemployment, and interest rates as $I(1)$.[5] For Germany, the interest rate in first differences shows a KPSS rejection of stationarity, possibly reflecting the structural shift from the long period of declining rates (2000–2021) to the sharp tightening cycle (2022–2024). However, the ADF-GLS provides supporting evidence for stationarity in first differences (p-value 0.040), and visual inspection (Figure 2) confirms the absence of persistent trends in the differenced series.

Log of house prices ($\ln h$) and log of construction costs ($\ln c$) present more complex integration properties. For house prices, the ADF-GLS test fails to reject the unit root null in levels even with a quadratic trend specification (p-values ≥ 0.749). Crucially, first differences also show high p-values with constant and linear trend for Germany, the Netherlands and Belgium, suggesting that a single differencing is insufficient to achieve stationarity. For France and Austria, the null can be rejected at the 5% level. However, the KPSS test

[5] There is various literature discussing stationarity characteristics of interest rates. For an early contribution based on nominal interest rates with unit root, as we use them, see for instance Rose, 1988.

suggests at least for France that the first difference is non-stationary. Only second differences display convincing evidence of stationarity, with p-values dropping to 0.024 or below for all countries. The KPSS test confirms this pattern. This provides strong statistical support for treating house prices as $I(2)$.

Log of construction costs ($\ln c$) show a less clear-cut pattern. The ADF-GLS fails to reject unit roots in levels with quadratic trend (p-values ≥ 0.161). An exception is Austria for which the null is rejected (p-value 0.012). First differences show mixed results: Germany and Netherlands display borderline evidence of stationarity (p-values 0.025), while France, Belgium, and Austria show weaker evidence (p-values ranging from 0.166 to 0.743). The KPSS test rejects stationarity for cost levels in all countries and for cost first differences in the Netherlands, but fails to reject for first differences in the remaining countries. Second differences are not rejected as stationary by KPSS across all countries.

Visual inspection (Figure 2) of the house price and construction cost series supports the $I(2)$ classification. Both variables exhibit pronounced trends with possible trend breaks around 2008 (Global Financial Crisis) and 2020–2021 (COVID-19 pandemic and subsequent supply chain disruptions). The sustained acceleration in house prices during 2014–2022 across all countries—documented by Ciocchetta et al. (2024) as contributing to overvaluation concerns—creates patterns consistent with $I(2)$ processes. Similarly, construction costs exhibit sharp acceleration during 2021–2023 driven by supply chain disruptions and energy price spikes in the wake of the Ukraine war. Given these institutional and visual considerations, combined with the strong statistical evidence for house prices and the borderline evidence for construction costs, we proceed by treating both variables as $I(2)$ and entering them in the ARDL models as first differences of logarithms ($\Delta \ln h_t$ and $\Delta \ln c_t$). This transformation ensures that the regressors are at most $I(1)$, satisfying the requirements of the ARDL framework. Rental prices ($\ln r$) also present an ambiguous pattern. The ADF-GLS test provides strong evidence against unit roots in levels for most countries (p-values ≤ 0.014 for Germany, Belgium, Austria, and borderline for Netherlands at 0.113), with France showing weaker evidence (p-value 0.181). First differences display mixed results. The KPSS test, however, rejects stationarity in levels for all countries, suggesting trend-stationarity rather than level-stationarity. In first differences, the KPSS fails to reject stationarity for most countries except France.

In summary, our unit root analysis provides statistical support for the following classifications: building permits, income, unemployment, and interest rates are treated as $I(1)$. House prices are clearly $I(2)$ while construction costs are $I(2)$ or at minimum exhibit strong persistence requiring differencing. Rents are $I(1)$ or possibly trend-stationary. These classifications align with the visual patterns documented in Figure 2. The ARDL framework accommodates this mixed integration structure, and subsequent cointegration tests will verify the existence of more or less stable long-run relationships.

3.4 ARDL-ECM specification

We model the log of building permits ($z_t \equiv \ln p_t$) as an ARDL consisting of levels except for house prices and construction costs with country-specific parameters and a common set of determinants $\boldsymbol{x_t}$ drawn from Section 2: real income per capita ($\ln y_t$), house price growth ($\Delta \ln h_t$), construction cost inflation ($\Delta \ln c_t$), rents ($\ln r_t$), the unemployment rate (u_t), and the mortgage rate on new loans (i_t).The pronounced long cycles visible in Figure 2, where building permits and their macroeconomic determinants exhibit persistent co-movements, motivate the application of cointegration analysis. The cointegration framework is specifically designed to capture such long-run equilibrium relationships. Construction activity, with its well-documented long cycles, represents a natural application of this framework (Burns, 1954).

Throughout this analysis, we will use the ARDL approach to cointegration (Pesaran et al., 2001). The ARDL approach is a flexible and robust method for estimating both the short- and long-run relationships between variables, even in the presence of mixed order of integration. It allows for the inclusion of both $I(0)$ and $I(1)$ variables in the same model, making it suitable for our analysis of building permits and their determinants. Recently, it has been applied by Ciocchetta et al. (2024) in their study on house price dynamics in Europe, and by Jonas et al. (2023) studying the responsiveness of new orders in the German main construction sector. All these studies implicitly restrict themselves to modelling one cointegration relation without discussing corresponding test results.

We consider the following ARDL(ℓ, q) model (notation similar to Shin et al. (2014))

$$z_t = \sum_{i=1}^{\ell} \phi_i z_{t-i} + \sum_{j=0}^{q} \boldsymbol{\theta}_j' \boldsymbol{x}_{t-j} + \epsilon_t \tag{1}$$

where $\boldsymbol{x}_t$ is a $k \times 1$ vector of multiple regressors, $\boldsymbol{\theta}_j$ is a $k \times 1$ vector of distributed-lag parameters, ϕ_i is the autoregressive parameterof lag i, and ϵ_t is an $i.i.d.$ process with zero mean and constant variance. For ease of exposition, we disregard the intercept and trend components in the model in Equation (1), even though these deterministics are included in our empirical models below. The ARDL(ℓ, q) model can be rewriten in the error correction (ECM) form as

$$\begin{aligned}\Delta z_t &= \rho z_{t-1} + \boldsymbol{\theta}' x_{t-1} + \sum_{i=1}^{\ell-1} \gamma_i \Delta z_{t-i} + \sum_{j=0}^{q-1} \boldsymbol{\pi}_j' \Delta \boldsymbol{x}_{t-j} + \epsilon_t \\ &= \rho \zeta_{t-1} + \sum_{i=1}^{\ell-1} \gamma_i \Delta z_{t-i} + \sum_{j=0}^{q-1} \boldsymbol{\pi}_j' \Delta \boldsymbol{x}_{t-j} + \epsilon_t \end{aligned} \tag{2}$$

where $\rho = \sum_{j=1}^{\ell} \phi_j - 1$, is the error correction coefficient, $\gamma_j = -\sum_{i=j+1}^{\ell} \phi_i$ for $j = 1, \ldots, \ell - 1$, $\boldsymbol{\theta} = \sum_{j=0}^{q} \boldsymbol{\theta_j}$, $\boldsymbol{\pi}_0 = \boldsymbol{\theta}_0$, and $\boldsymbol{\pi}_j = -\sum_{i=j+1}^{q} \boldsymbol{\theta}_i$ for $j = 1, \ldots, q-1$. Furthermore, $\zeta_{t-1} = z_{t-1} - \boldsymbol{\beta}' \boldsymbol{x}_t$ is the error correction term where $\boldsymbol{\beta} = -\boldsymbol{\theta}/\rho$, for $\rho \neq 0$, are the associated long-run parameters.

The conditional ECM rests on the assumption that the level of z_t does not feed back into the marginal process for $\boldsymbol{x}_t$. Under this assumption, $\boldsymbol{x}_t$ is long-run forcing for z_t, and the analysis is restricted to at most one conditional level relationship between z_t and $\boldsymbol{x}_t$, irrespective of the integration order of $\boldsymbol{x}_t$ (Pesaran et al., 2001, Assumption 3, p. 293).

We follow the conditional modeling approach in which contemporaneous changes $\Delta \boldsymbol{x}_t$ are included as regressors, potentially correcting for the weak endogeneity of any $I(1)$ explanatory variables (Shin et al., 2014). Under the implicitly assumed weak exogeneity and given an appropriate lag structure and absence of residual series correlation, standard OLS provides consistent estimates of both the long-run parameter β and the short-run dynamic parameters (Shin et al., 2014).

We implement two complementary cointegration tests tailored to ARDL-ECM. The BDM test (Banerjee et al., 1998) is based on the t-statistic for the exclusion of the lagged dependent variable z_{t-1} in the ECM, formally, $\rho = 0$ against $\rho < 0$. The PSS test (Pesaran et al., 2001) is based on the F-test of the joint null that $\rho = 0$ and each of the $\boldsymbol{\theta}$-values is zero in (2). The two tests are complementary because they guard against distinct failure modes: a significant PSS F-statistic with an insignificant BDM t-statistic signals a degenerate case in which the regressors enter the long-run relationship but there is no error-correction restoring force, implying no genuine cointegration (McNown et al., 2018). For both tests, we apply bootstrap methods to eliminate inconclusive inference of the asymptotic bounds procedure and to improve finite-sample size properties (McNown et al., 2018; Bertelli et al., 2022).[6] Our application of the ARDL approach with cointegration testing follows established best practices in the recent housing literature. Ciocchetta et al. (2024) also apply a single-equation ARDL model for analyzing house price determinants across European countries. We also rely on the assumption of weak exogeneity for all regressors. To account for outliers, we add country-specific impulse dummies. This helps to obtain stable and statistically sound specifications as reported by the standard battery of tests on serial correlation, heteroskedasticity and functional form.

4 Empirical results

4.1 Triangular specification design

We adopt a triangular (nested) specification strategy to balance parsimony, identification, and cross–country comparability. Starting from a minimal demand anchor, we progressively add variables following the classification established in Table 1, assessing at each step whether a long-run relation can be established via cointegration tests, while tracking sign stability and magnitudes of the long-run coefficients. The design transparently reveals mediation effects, for instance, income effects absorbed by prices, and mitigates multicollinearity.

Concretely, for each country we estimate six nested ARDL–ECM models for $\ln p_t$. The ordering proceeds from demand-side to financing cost to supply-side variables, and finally to asset market signals: (i) real income per capita, (ii) the mortgage rate, (iii) growth of construction costs, (iv) the unemployment rate, (v) growth of house prices, and (vi) rents while omitting income to test for mediation by asset-market variables. This sequencing is motivated by the empirical observation, emphasized by Duca et al. (2021), that purely supply-side specifications struggle empirically, while combining demand-side and supply-side indicators improves model performance. The final specification—replacing income with rents—explicitly tests whether rental profitability mediates the income-to-construction channel, as suggested by the theoretical framework in Section 2.

[6] We employ a wild-bootstrap procedure for computing the standard errors of long-run multipliers as well for computing the confidence intervals of the dynamic multipliers to a permanent change in some regressor.

The following overview lists the regressor vector for each specification, where we account for the $I(2)$ properties of construction costs and house prices by entering them as first differences of their logarithms:

(1) $\boldsymbol{x}_t = (\ln y_t)$

(2) $\boldsymbol{x}_t = (\ln y_t, i_t)$

(3) $\boldsymbol{x}_t = (\ln y_t, i_t, \Delta \ln c)$

(4) $\boldsymbol{x}_t = (\ln y_t, i_t, \Delta \ln c, u_t)$

(5) $\boldsymbol{x}_t = (\ln y_t, i_t, \Delta \ln c, u_t, \Delta \ln h_t)$

(6) $\boldsymbol{x}_t = (i_t, \Delta \ln c, u_t, \Delta \ln h_t, \ln r_t)$

Each specification includes an intercept, linear trend and country-specific impulse dummies to account for outliers.

4.2 Summary of Cointegration Test Results

As the results between both tests are qualitatively similar, we focus our discussion on the PSS bounds test results reported in Table 7. The BDM test results are provided in the Appendix in Table A1. We conclude in favor of cointegration where the PSS test rejects the null hypothesis at the 5% or 10% significance level.

– Place Table 7 here. –

A clear pattern emerges across the countries. For Germany, cointegration is borderline when adding the mortgage rate to income (Model 2: $p = 0.085$), but becomes significant at the 5% level once construction cost dynamics are included (Model 3: $p = 0.032$). The evidence becomes very strong when unemployment (Model 4) or asset-market variables are included (Models 4-6: $p \leq 0.003$). This suggests that while income and interest rates are crucial, supply-side factors (costs) and cyclical controls are necessary to fully capture the long-run equilibrium of German building permits.

France and Belgium display the most robust evidence. For France, adding the mortgage rate to income is already sufficient to establish a strong long-run relationship (Model 2: $p < 0.001$), with robust rejections of the null in all richer specifications. Belgium displays pervasive evidence of cointegration across all six specifications ($p < 0.001$ throughout), implying that income alone (Model 1) can anchor a long-run relation.

For the Netherlands, we find strong evidence of cointegration in Models 1 through 5 ($p \leq 0.008$), indicating that combinations of income, interest rates, and house price appreciation form stable long-run relationships with permits. In the final specification including rents

(Model 6), the PSS test indicates cointegration at the 10% level ($p = 0.085$). However, the BDM test provides stronger support for Model 6 ($p = 0.019$, see Appendix), rejecting the null of no cointegration at the 5% level. Given the robust rejection across Models 1–5 and the supporting BDM evidence, we consider the evidence for a Dutch long-run relationship to be generally robust, though slightly weaker when rents are included.

By contrast, Austria does not pass the cointegration test in any specification. The minimum p-value observed is 0.239 (Model 4), with other specifications showing $p > 0.35$. Unlike the other countries in our sample, we find no statistical evidence of a stable long-run relationship between building permits and macroeconomic fundamentals for Austria. Consequently, the subsequent discussion of long-run coefficients focuses on Germany, France, the Netherlands, and Belgium. However, we will report dynamic multipliers for all countries including Austria further below. For Austria, estimation will then be restricted to first-differences given no long-run relationship.

4.3 Long-run coefficient estimates

This section presents the bootstrapped long-run coefficient estimates for the ARDL models. As established in Section 4.2, we restrict our detailed interpretation to Germany, France, the Netherlands, and Belgium. Concerning the regressors, it is crucial to note the interpretation of construction costs and house prices. Due to the $I(2)$ nature of the level series, these variables enter the model as first differences of logarithms ($\Delta \ln c_t$ and $\Delta \ln h_t$). Consequently, their coefficients represent semi-elasticities with respect to the growth rate. A positive coefficient for $\Delta \ln h_t$ implies that an acceleration in house price growth (speculative momentum) stimulates permits.

Several important patterns regarding the long-run effects emerge from the cross-country comparison, summarized in Table 8.

– Place Table 8 here. –

Germany Table A2 reports detailed estimation results. Model fit is consistently high ($\bar{R}^2 \approx 0.92$) with clean model diagnostics. The error correction coefficient (ρ) strengthens from -0.04 (Model 1) to -0.14 (Models 4–6), suggesting that supply-side and labor market controls are necessary to anchor a well-identified equilibrium. Interest rate semi-elasticities are negative and highly significant (1% level) in Models 2–5, ranging from -0.20 to -0.26. This implies a long-run reduction in permits of 20–26% per percentage point increase in rates. Significance vanishes in Model 6 when rents are included. Income elasticities are large and positive in Models 4–5 (5.4) but borderline insignificant (t-statistics ≈ 1.5). Unemployment reduces permits by 8–9% per percentage point (Models 4–6, 1% level) which is slightly lower than the range reported by Jonas et al. (2023). Construction cost inflation shows negative point estimates across Models 3–6 (-44 to -16) but remains statistically insignificant. House price appreciation is insignificant in Models 5–6. Rental levels (Model 6) show a significant positive elasticity of 6.1 (5% level).

France Table A3 reports detailed results. Model fit is robust ($\bar{R}^2 \approx 0.90$), though Models 3–5 show some evidence of first-order serial correlation. The speed of adjustment is notably faster than in Germany ($\rho \approx -0.22$ to -0.24 in Models 2–5), implying that approximately one-quarter of any disequilibrium is corrected each month. Income elasticities are robustly significant at the 1% level (4.9 to 5.3, Models 2–5). Interest rate semi-elasticities are negative and significant at the 1% level in Models 2–5 (-0.14 to -0.23), but insignificant in Model 6. House price appreciation is positive and significant (15.5 in Model 5, 33.3 in Model 6). Construction cost inflation shows significant positive coefficient in Models 3–5 (11.3 to 13.8). Rents (Model 6) are positive and highly significant (2.72, 1% level). Unemployment is generally insignificant, except for a counter-intuitive positive coefficient in Model 5.

Netherlands Table A4 reports detailed results. Model fit ranges from $\bar{R}^2 = 0.77$ to 0.79 with no significant serial correlation. Error correction ranges from -0.18 (Model 1) to -0.28 (Model 4), with Model 6 showing markedly slower adjustment ($\rho = -0.19$). Income elasticities are exceptionally high and robustly significant (8.3 to 10.6, Models 1–5). Interest rate effects are weak and unstable: borderline significant only in Model 2 (-0.12), with significance vanishing in richer specifications. House price appreciation is large and highly significant (31.0 in Model 5, 60.9 in Model 6). Construction costs show positive and significant coefficients in Models 3–5 (12.1 to 18.3). Unemployment is insignificant except for a borderline negative effect in Model 6. Rents show a significant negative coefficient (-5.3, 5% level).

Belgium Table A5 reports detailed results. Model fit is noticeably lower than for other countries ($\bar{R}^2 = 0.55$–0.57), but Belgium exhibits the fastest error correction ($\rho = -0.20$ to -0.39). Interest rate coefficients are consistently negative (-0.04 to -0.09) with varying significance: Model 4 achieves 1% significance, while other specifications show weaker effects. Income and unemployment are insignificant across all specifications. House price appreciation shows a borderline positive coefficient in Model 5 only (18.7, approaching 10%). Construction costs show mixed signs and remain insignificant. Rents are positive but insignificant.

4.4 Cross-country patterns and implications

Building on the country-specific estimation results summarized above, we now compare findings across countries, organized by transmission channel (see Table 1 for the theoretical predictions). Furthermore, we relate our findings to institutional differences documented in the comparative housing literature.

Interest rate effects (financing costs.) We find consistently negative long-run effects in all four cointegrated countries, confirming the theoretical predictions (Table 1) that financing costs dampen construction activity. However, the magnitude and robustness of the effect are heterogeneous: Germany and France exhibit strong and robust sensitivity, Belgium shows significant effects only in selected specifications, and the Netherlands displays significant effects only in the most parsimonious model, with significance vanishing once house prices are included.

This cross-country heterogeneity aligns with differences in housing finance institutions. While higher borrowing costs generally reduce credit access and increase debt service burdens (Paz-Pardo et al., 2024), the strength of this transmission depends on institutional intermediation. The Netherlands' Mortgage Interest Deduction, which allows homeowners to deduct mortgage interest payments from taxable income, and the National Mortgage Guarantee, which reduces risk premia, both dampen sensitivity to gross interest rate changes (Gielens, 2010). Joebges et al. (2015) provide a structural explanation: the Netherlands belongs to the high credit-to-GDP group, where monetary policy affects housing markets primarily through credit-amplified asset price dynamics rather than through direct cost-of-capital effects on construction. In a similar vein, Tarne and Bezemer (2025) reveal trade-offs between shocks to the housing supply, to interest rates and to banks' loan-to-value settings by estimating the corresponding effects on Dutch house prices. Melecky and Paksi (2024) confirm this asset-price channel, finding that long-term interest rates have a significant negative effect on housing prices across 15 European countries. Our finding that Dutch interest rate significance vanishes once house price inflation is included is consistent with both results: the house price variable absorbs the interest rate channel because asset price transmission is the dominant pathway through which monetary policy operates in high-credit environments.

Germany, lacking broad-based mortgage subsidies, exhibits the strongest direct transmission from financing costs to construction. This is consistent with Shida (2022), who documents that the interest rate channel is stronger in countries with more liberalized mortgage markets. The heterogeneity also reflects differences in credit market structure documented by Kelly et al. (2019): in boom-bust countries like the Netherlands, pre-2008 credit loosening created borrower populations more sensitive to asset price movements than to financing costs, while stable credit standards in Germany and France maintained a direct link between borrowing costs and construction decisions.[7]

Income effects. We identify a clear dichotomy between a demand-driven regime and other influences. France and the Netherlands exhibit large and robustly significant income elasticities, while Germany and Belgium show weak or insignificant effects.

The strength of income transmission in France and the Netherlands reflects institutional amplifiers. France's DALO Act (Droit au Logement Opposable) of 2007 grants individuals an enforceable right to housing and obligates the state to provide accommodation (Loison-Leruste and Quilgars, 2009), institutionalizing the link between income growth and housing demand. Caldera and Johansson (2013) estimate that the Netherlands has the lowest long-run price elasticity of housing supply among the five countries in our sample, implying that income-driven demand shocks translate disproportionately into price adjustments rather than quantity responses—which may amplify the apparent income elasticity of permits, as

[7]Recent market developments underscore these patterns. In Germany, interest rates rose from a historic low of 1.16% in late 2020 to 4.05% by December 2023, contributing to a halving of building permits. In France, rates rose from 1.1% in December 2021 to 3.6% over a similar period, contributing to an 11% decline in house prices and a nearly 30% drop in building permits. Belgium experienced the sharpest increase to 3.9% by December 2023—the highest level since September 2011—accompanied by a roughly 50% decline in permits from mid-2021 to August 2023.

income growth raises demand that the rigid supply side cannot accommodate.[8]

The weakness of income effects in Germany likely reflects mediation through the rental market. As documented in the country-specific results, replacing income with rents in Model 6 yields a highly significant positive rent coefficient, suggesting that income transmits to construction via rental demand rather than directly through owner-occupier housing demand. Additionally, Germany's stringent land-use regulations and lengthy approval processes—evidenced by the persistent failure to meet the 400,000-unit annual construction target (with recently only 294,400 (2023) and 251,900 (2024) completions)—may prevent income shocks from translating into quantity adjustments (Cavalleri et al., 2019). For Belgium, the insignificance of income may reflect the substantial role of tax incentives in housing policy, which can decouple construction decisions from contemporaneous income dynamics. The weaker model performance may also partly reflect Belgium's fragmented housing policy landscape, where regional governments (Flanders, Wallonia, Brussels) maintain substantial autonomy over housing regulations.

The shift in borrower composition documented by Kelly et al. (2019) provides additional insight. In Germany and Austria, main-residence ownership has decreased for the youngest cohort (born after 1980) while increasing for older cohorts. In France, and the Netherlands, by contrast, young households have *increased* their ownership rates. This divergence suggests that where younger, more income-constrained households dominate the marginal buyer pool, aggregate income elasticities are higher; where older, wealthier households drive demand, income effects are attenuated.

House price appreciation (speculative channel). House price dynamics emerge as a powerful driver in three of the four cointegrated countries but are entirely absent in Germany, reinforcing the regime distinction. The cross-country pattern maps closely onto the institutional classifications established by Joebges et al. (2015) and Kelly et al. (2019).

Germany is the only country where house price appreciation is insignificant, indicating that developers respond to cost and financing constraints rather than speculative momentum. This aligns with Germany's classification as a non-boom-bust country with stable credit standards throughout the 2000s (Kelly et al., 2019) and the absence of housing bub-

[8] More broadly, Caldera and Johansson (2013) show that the number of days required to obtain a building permit is negatively correlated with supply responsiveness across OECD countries, a relationship that holds both across countries and across US cities.

bles documented by Joebges et al. (2015) for the 1990–2012 period—an assessment confirmed by Feld et al. (2026) for the subsequent decade. Survey evidence from Kindermann et al. (2021) provides a micro-foundation: German homeowners systematically under-predict house price growth during booms, attributing increases to recovery from financial distress (expecting mean reversion) rather than to fundamental rent growth (expecting momentum). Without extrapolative expectations, no self-reinforcing feedback loop between house prices and construction can emerge. However, our insample analysis should not be read to suggest that no future price bubbles will be originated in the German housing market and that the identified regime will be stable over time.[9]

A complementary structural explanation for country-specific differences lies in the composition of the building sector. As discussed in Section 2, commercial developers respond directly to house price profitability signals, while self-building households do not. Germany's comparatively large self-build share and France's and the Netherlands' more developer-dominated markets are therefore consistent with the observed pattern: significant house price coefficients where developer-driven construction dominates, and insignificant where it does not.

The Netherlands shows the strongest house price response, consistent with its classification in the high credit-to-GDP group (Joebges et al., 2015) and as a boom-bust market (Kelly et al., 2019; Tarne and Bezemer, 2025). Pre-crisis credit loosening—with nearly half of all mortgages originated at loan-to-value (LTV) ratios of 90% or higher—created strong linkages between price expectations and construction. Although credit standards have since tightened, the speculative channel appears structurally embedded, as evidenced by our finding that house price appreciation dominates interest rate effects in richer specifications. The "Young Buys Old" program, designed to redirect demand toward existing stock, reflects policy recognition of these dynamics (Boelhouwer, 2017).

France arrives at a similar outcome through a structurally different channel. Despite belonging to the low credit-to-GDP group alongside Germany, France experienced two housing bubbles between 1985 and 2012 (Joebges et al., 2015). This means French speculative dynamics cannot be attributed to credit amplification. Instead, institutional investment incentives—the Pinel Law encouraging buy-to-let construction (Lévy, 2022) and the DALO

[9]Note that Deutsche Bundesbank has regularly warned of significant (urban) overvaluations in its Financial Stability Reviews.

Act institutionalizing housing demand—appear to substitute for the credit channel, generating speculation through policy design rather than financial depth.

Belgium occupies an intermediate position. Joebges et al. (2015) classify Belgium as a non-bubble country despite substantial house price appreciation (over 70% between 2000 and 2021). The borderline significance of house price effects in our estimates (Model 5 only, approaching 10%) is consistent with this intermediate classification: speculative dynamics are possibly present but structurally weaker than in France or the Netherlands.

Construction cost effects. The theoretically expected negative sign (Table 1) emerges only in Germany—and, as shown below, in Austria's dynamic multiplier analysis—but without statistical significance in the long-run estimates. France shows significant positive coefficients, while the Netherlands and Belgium display positive effects with varying significance. This cross-country heterogeneity in a key supply-side variable warrants detailed discussion.

Germany's unique (though insignificant) negative cost sign may reflect the multi-dimensional cost pressures documented by Feld et al. (2026): construction prices stood 48.5% above five-year-ago levels as of 2025q3, driven by sharply rising energy and material (for concrete and cement) prices, higher construction wages, and stricter energy efficiency regulations under the Building Energy Act (GEG) which came into effect in 2024 and introduced stringent standards for heating systems and insulation. However, the statistical insignificance of the effect, despite these well-documented cost pressures, is consistent with the levels-versus-changes distinction identified by Lerbs (2014). Estimating German county-level permit equations that include construction costs in both levels and first differences simultaneously, he finds that cost *levels* carry the expected negative sign while cost *changes* show a significant positive coefficient, attributed to anticipatory investment: developers extrapolating rising costs accelerate projects to build before further escalation. Since our ARDL specification captures cost changes ($\Delta \ln c$), these two effects partially offset each other, attenuating the estimated coefficient toward zero. In Germany, the negative cost-level effect may still dominate.

For France and the Netherlands, the positive cost coefficients may be attributed to measurement error (DiPasquale, 1999) or demand-pull dynamics. However, two alternative explanations deserve considerations. First, the dominance of the anticipatory change effect documented by Lerbs (2014): during the sharp cost acceleration of 2021–2023, developers facing rapidly rising material and energy prices had strong incentives to advance projects before further escalation. Second, an indirect channel operating through house prices: Melecky and Paksi (2024) find that construction costs have a significant positive effect on housing prices across 15 European countries. If rising costs push up house prices, and house prices drive construction—as we find for France and the Netherlands—the positive cost coefficient in our supply equations may partly reflect this price-mediated transmission rather than measurement failure. Jonas et al. (2023) similarly find a positive cost coefficient for German new construction orders but do not control for house prices, which may explain why the indirect price channel is not disentangled in their estimates.

Our results can be situated in the broader cross-country literature. Caldera and Johansson (2013), estimating long-run supply equations for 21 OECD countries using residential investment, report positive and significant cost coefficients for Belgium and Germany, and a positive though insignificant coefficient for France. That the puzzle persists across different dependent variables (residential investment versus building permits), estimation frameworks, and sample periods (1980s–2000s versus 2000–2024) reinforces the interpretation that it reflects a fundamental identification challenge. At least three factors may account for the differences in the sign of the cost coefficient in Germany: (i) their sample was more strongly influenced by the divergent post-reunification dynamics in eastern and western Germany (Knetsch, 2010); (ii) our sample is less affected by the macroeconomic wage restraint of the early 2000s (Dustmann et al., 2014), which also dampened construction sector costs; and (iii) unlike theirs, our sample includes the material cost inflation driven by COVID-19 supply chain disruptions (Moosavi et al., 2022). These structural differences probably explain the different outcome.

Unemployment effects. Table 1 predicts a negative effect of unemployment on construction through reduced income security and tighter credit access. However, we find robust direct effects only in Germany, where unemployment is significant at the 1% level across Models 4–6. The Netherlands shows a borderline negative effect in Model 6 only, Belgium displays no significant effects, and France exhibits a counter-intuitive positive coefficient in Model 5. The absence of robust unemployment effects in three of the four cointegrated countries suggests that labor market conditions influence construction primarily through indirect channels—via income (captured by $\ln y$) or credit availability (captured by i)—rather than through a direct transmission mechanism. The Netherlands illustrates this pattern: despite a sustained labor market tightening (unemployment fell from 8.7% in March 2014 to 3.6% by late 2023), the direct unemployment effect remains insignificant in most specifications, suggesting that the improvement transmitted to construction via rising incomes and house prices rather than through the unemployment rate per se. This interpretation is consistent with the observation that unemployment effects strengthen in Germany precisely when income is included in the specification (Models 4–6), suggesting complementarity rather than substitutability between the two demand-side variables.

Rental market effects. The rental channel (Table 1) operates selectively across countries, with sign and significance varying in ways that illuminate the regime distinction. Germany and France exhibit significant positive rent elasticities, while the Netherlands displays a significant negative coefficient. Belgium shows a positive but insignificant effect.

The positive rent effects in Germany and France are consistent with the buy-to-let channel documented by Paz-Pardo et al. (2024): higher rents signal profitable investment opportunities that stimulate construction. At the same time, higher rents might stimulate construction activity of future owner occupiers to avoid rental payment. In Germany, with its relatively low homeownership rate, rent regulation—including the Mietpreisbremse (rent brake) and Kappungsgrenze (rent cap)—makes rental income more predictable and thus a more reliable signal for developers than volatile house prices. Recent policy initiatives such as the "Neue Wohngemeinnützigkeit," reinstated in January 2025 to incentivize non-profit housing provision through tax benefits (BMWS, 2024), further underscore the policy emphasis on stabilizing rental markets. Kindermann et al. (2021) provide micro-foundations for this channel: real estate investors, including non-corporate landlords, who own rental properties, form expectations similar to renters, confirming that direct exposure to the rental market improves information quality and investment decisions. In France, the Pinel Law offers substantial tax reductions for investment in new rental properties (Lévy, 2022), explicitly incentivizing the buy-to-let channel. With over 40% of French renter households residing in public housing, private rental construction may be particularly responsive to profitability signals in the remaining market segment.

The Netherlands' negative rent coefficient presents a mirror image of the German case. In Germany, where the credit-to-GDP ratio is low and the rental market well-developed, rents serve as the primary investment signal (positive coefficient). In the Netherlands, where credit depth and securitization levels are high (Joebges et al., 2015), capital gain expectations dominate the investment calculus, relegating rents to a counter-cyclical indicator: rising rents may signal a tightening ownership market—with households pushed into renting because they cannot afford to buy—which coincides with reduced construction activity driven by weakening price momentum.

Austria: The non-cointegrated case. Austria's failure to establish cointegration across all specifications may itself be institutionally informative. Like Germany, Austria exhibits low homeownership rates and belongs to the non-boom-bust group with stable credit standards throughout the 2000s (Kelly et al., 2019). The dominant non-profit housing sector—where limited-profit housing associations (Genossenschaftswohnungen) and municipalities (Gemeindewohnungen) contribute to over half of multi-apartment construction, especially in Vienna—operates largely outside market dynamics (IMF, 2024). Independent evidence corroborates this structural decoupling. Caldera and Johansson (2013), estimating supply equations for OECD countries within a stock-flow error correction framework over the 1980s to mid-2000s, find that Austria is the only country in their sample where the error correction term in the investment equation carries a positive sign, "implying basically no adjustment" towards a long-run (market) equilibrium.

Two institutional layers explain why standard macroeconomic channels fail to anchor a long-run relationship. First, as documented by Joebges et al. (2015), housing assistance schemes subsidize construction and renovation through interest rate subsidies rather than tax-based systems, decoupling the effective cost of borrowing from the market rate. Strong tenant protection and limited tax deductibility of mortgage interest further discourage speculative buys-to-let investment. Second, macroprudential credit regulation (KIM-Verordnung) requiring 20% equity for housing loans and capping debt service at 40% of household income constrains not only the costs but also access to credit independently of monetary policy conditions (Luckert, 2026). Together, these layers create a bifurcated market where private construction follows policy priorities rather than market fundamentals.

Summary: Two regimes in the long-run. The cross-country comparison reveals two distinct regimes shaped by institutional differences in credit markets, rental structures, and housing policy. Germany operates as a "fundamental and supply-side driven" regime and is the only country where both the financing cost and supply-side channels identified in Table 1 operate simultaneously. Developers respond to observable rental income rather than potentially noisy house price signals—a pattern consistent with the information structure documented by Kindermann et al. (2021).

France and the Netherlands display "speculative and demand-side driven" regimes where income and house price momentum dominate, though through structurally different channels: credit amplification in the Netherlands versus institutionally created investment incentives in France. For both, house price bubbles are well documented in the literature.

Belgium occupies an intermediate position consistent with its classification as a non-bubble country (Joebges et al., 2015), with demand-driven dynamics to some extent but structurally weaker model fit, possibly reflecting its fragmented regional policy landscape.

4.5 Dynamic Multiplier Results

This section analyzes the dynamic responses of building permits to permanent changes in key macroeconomic variables (Table 1) over a 16-month horizon, adding to the previous long-run analysis the effects from short-run dynamics. For Germany, France, the Netherlands, and Belgium, we report multipliers from the cointegration specification established in Section 4.3. For Austria, where no cointegrating relationship could be established, we report results from a first-difference specification; these capture short- to medium-run dynamics in permit *growth rates* and do not have a long-run equilibrium interpretation. The dynamic multiplier analysis complements the long-run coefficient estimates by revealing adjustment speeds and potential overshooting or delayed transmission. Tables 9 and 10 provide an overview of the key findings across all shocks.

– Place Table 9 here. –

– Place Table 10 here. –

Interest rate responses. Figure 3 depicts the dynamic response to a one percentage point permanent increase in interest rates. The results reinforce the regime distinction: the "fundamental and supply-side driven" regime exhibits the direct financing cost transmission, while the "speculative and demand-side driven" regime countries show weaker responses.

Germany and France exhibit the most robust patterns: Models 3–5 show significantly negative effects emerging after 4–6 months, with multipliers converging to approximately -0.15 (-0.2) for Germany (France) after 16 months. In Model 6, where rents absorb part of the financing cost channel, the effect loses significance within the 16 month horizon.

The Netherlands presents markedly weaker short- to medium-run evidence, consistent with a regime where monetary policy transmits primarily through asset prices rather than direct cost-of-capital effects. Models 4–6 show no significant effects throughout the horizon. Only Model 3 displays borderline significance after 10 months. Point estimates for Models 3, 4, and 6 converge to similar negative values after 16 months, while Model 5 yields implausibly positive point estimates throughout. For Belgium, Models 4 and 6 produce significantly negative effects after approximately 8 months, converging to a multiplier of -0.1. The remaining models yield similar point estimates but without statistical significance.

Austria, estimated in first differences, also displays negative responses consistent with the direct financing cost channel observed in Germany. Model 4 achieves significance after 3 months, Model 3 after approximately 7 months. Mean responses are negative across all models, though Models 5 and 6 do not reach statistical significance.

Income responses. Figure 4 depicts the dynamic response to a permanent income change. France and the Netherlands exhibit strong and significant effects across Models 3–5. Both countries display initial overshooting in the first month, followed by a brief correction, before stabilizing at significant positive multipliers. France reaches approximately 3.0 after 4 months and 5.0 at 16 months, while the Netherlands stabilizes at approximately 4.0 after 2 months and 9.0 after 16 months.

Germany and Belgium show weaker income transmission, consistent with their classification outside the demand-driven regimes. For Germany, Models 4 and 5 display marginally significant positive multipliers emerging after approximately 4 months, reaching about 4.4 after 16 months; Model 3 shows no significant effects. Belgium exhibits a significant positive effect only for Model 4, emerging after about 9 months with a multiplier approaching 2.0 after 16 months. The remaining Belgian models show high uncertainty throughout.

Austria presents negative point estimates in the first-difference specification across all three models during the initial 8 months, with only borderline significance during the first 2 months. This counter-intuitive pattern likely reflects the limitations of the first-difference transformation, which removes the levels information necessary to anchor income effects.

Unemployment responses. Unemployment effects (Figure 5) display pronounced cross-country heterogeneity. Germany shows a counter-intuitive immediate positive reaction for Models 4–6 lasting only few months, before the effect turns negative and significant after approximately 11–13 months, with similar responses across all three specifications. This delayed adjustment—the longest among all determinants in the German "fundamental and supply-side driven" regime—is consistent with the complementarity between income and unemployment documented in the long-run analysis: unemployment shocks require time to erode income security and tighten credit access before affecting construction planning.

France displays persistently counter-intuitive positive responses across the models, though with high uncertainty. The mean effect decreases towards zero for Model 6 but remains positive for Models 4–5. Given the insignificance of unemployment in the long-run estimates, these short-run patterns likely reflect specification noise rather than a genuine transmission—consistent with a "speculative and demand-side driven" regime where asset price and income dynamics absorb labor market signals. The Netherlands shows an immediate negative and significant effect of approximately -0.10 that dissipates after one month. Only Model 6 exhibits a lasting negative effect, and Model 5 produces a counter-intuitive positive effect after approximately 7 months. Belgium displays a similar pattern, with only Model 6 indicating borderline significance for a negative mean multiplier after 3 months.

Austria exhibits theoretically consistent unemployment effects in the first-difference specification paralleling the direct transmission observed in Germany. Model 4 shows significantly negative effects after 4 months (multiplier approximately -0.03), Model 5 reaches borderline significance after 5 months, while Model 6 does not achieve lasting significance.

House price responses. House price effects (Figure 6) sharply differentiate the two regimes identified above. Germany shows no significant effects for either Model 5 or Model 6 at any horizon, reinforcing its classification as a "fundamental and supply-side driven" regime where speculative price momentum plays no role in construction decisions.

France, the Netherlands, and Belgium all display significant positive responses consistent with the "speculative and demand-side driven" regime. France shows significantly positive effects for both models after 3–4 months, with Model 6 exhibiting stronger dynamics: multipliers of approximately 10.0 at 4 months and 28.0 at 16 months. The Netherlands displays comparable timing but larger magnitudes, with multipliers reaching approximately 15.0 at 4 months and exceeding 40.0 at 16 months—the strongest house price transmission among all countries. This is consistent with the credit-amplified speculative channel documented in the long-run estimates. Belgium shows significant effects only for Model 5, emerging after 3 months with a multiplier around 10.0 at 4 months and approaching 20.0 at 16 months.

Austria, in the first-difference specification, shows immediate and lasting positive effects for both models, with multipliers between 1.0 and 2.0 that remain significant throughout most of the 16 month horizon. This suggests that Austrian permit growth responds to some extent positively to house price acceleration, but as a reminder it is worth to mention that no stable long-run relationship could be established.

Building cost responses. Building cost effects (Figure 7) reinforce the regime distinction observed for other determinants. Germany and Austria, the two countries where the "fundamental constraint" logic applies, are the only ones displaying the theoretically expected negative response. For Germany, Models 3–6 yield comparable negative multipliers throughout all horizons, though uncertainty is large and significance remains borderline for modest specifications and horizons. Austria shows more robust evidence: significantly negative effects emerge for all models after approximately 9–10 months, making it the clearest case of theoretically consistent cost transmission in our sample.

France exhibits significantly positive effects after one month for Models 3–6, consistent with the anticipatory investment interpretation developed in the long-run analysis: developers accelerate projects in response to rising costs. The Netherlands shows similar short-run positive dynamics across Models 3–6, with borderline significance for Models 3–5 and insignificant effects for Model 6. Short-run patterns in both countries mirror the "speculative and demand-side driven" regime, where cost increases coincide with—and may partly operate through—rising house prices rather than constraining construction directly. Belgium produces negative mean responses for Models 3–5 consistent with supply-side theory, but the effect remain insignificant at all horizons. Model 6 shows positive but insignificant effects.

Rent responses. Rent effects (Figure 8) operate with the longest lags among all determinants and exhibit high uncertainty, consistent with rents being a slow-moving signal that requires sustained changes before developers adjust planning decisions. Germany shows high initial uncertainty, but the effect becomes significantly positive after 8 months (multiplier approximately 2.0) and steadily increases to approximately 5.0 at 16 months. This gradual build-up is consistent with the "fundamental and supply-side driven" regime, where developers rely on observable rental income as a profitability signal and self-building households, in light of rising rents, are encouraged to switch from tenant to homeownership.

France displays large fluctuation during the first 6 months, with significance emerging after 8 months (multiplier approximately 1.5). The slow adjustment contrasts with the rapid transmission of house price and income shocks documented above. The Netherlands shows initial dynamics similar to France, but the effect never achieves sustained significance. The mean effect turns negative after 10 months but remains statistically indistinguishable from zero. This weak dynamic response—despite a significant negative long-run coefficient—suggests that the long-run rental effect operates primarily through very slow stock adjustment rather than through the 16-month horizon captured here. Belgium exhibits a positive mean effect (multiplier approximately 1.5) after 4 months, but significance remains borderline throughout the horizon.

Austria displays a brief window of significance in the first-difference specification: a negative effect between months 6 and 8 (multiplier approximately -2.0), but effects are otherwise insignificant. This transitory response likely reflects specification noise rather than a genuine rental transmission channel, consistent with the long-run analysis.

Summary. The dynamic multiplier analysis confirms the two-regime typology established in the long-run estimates and reveals a consistent hierarchy of adjustment speeds across determinants. Income and house price effects materialize fastest (within 2–4 months), consistent with asset price signals transmitting rapidly through financial markets. Interest rate effects require 4–8 months, reflecting the slower pass-through of financing costs to developer project decisions. Unemployment effects display the longest lags (11–13 months in Germany), consistent with the indirect transmission through income security and credit access documented in the long-run analysis. Rent effects operate on a similar timescale (8+ months), reinforcing their role as a slow-moving but persistent investment signal.

This hierarchy has direct implications for policy timing. Demand-side stimulus through asset markets (e.g., house purchase subsidies) can be expected to affect construction within one quarter, whereas financing cost interventions (e.g., subsidized mortgage programs) require two quarters, and labor market improvements may take over a year to translate into construction activity. The "speculative and demand-side driven" regimes (France, the Netherlands) respond most strongly and quickly to asset price and income shocks, while the "fundamental and supply-side driven" regime (Germany) displays slower but broader-based responses across multiple transmission channels mentioned above.

5 Conclusion

This paper investigates the macroeconomic determinants of residential building permits in Germany, France, the Netherlands, Belgium, and Austria using monthly data from 2000 to 2024. Applying ARDL cointegration methods with bootstrap inference across six nested specifications per country, we test for stable long-run relationships between permits and a comprehensive set of determinants including income, interest rates, construction costs, unemployment, house prices, and rents.

Our results reveal two distinct regimes shaped by institutional differences in mortgage markets, housing policy, and rental market structure. Germany operates as a "fundamental and supply-side driven" regime: construction responds to financing costs, unemployment, and rental income, while house price appreciation hardly plays a role. Germany is the only country where both the financing cost and supply-side channels operate simultaneously, and the only one displaying the theoretically expected negative construction cost sign—though without statistical significance for the long-run multiplier. France and the Netherlands display "speculative and demand-side driven" regimes where income elasticities and house price momentum dominate, albeit through structurally different channels: credit amplification in the Netherlands (Joebges et al., 2015) versus institutionally created investment incentives in France (Lévy, 2022). Belgium occupies an intermediate position with weaker statistical evidence overall. Austria fails to establish cointegration in any specification, a finding we attribute to its dominant non-profit housing sector that insulate construction from market-based transmission channels (IMF, 2024).

The counter-intuitive positive cost coefficients observed for France and the Netherlands are consistent with the anticipatory investment behavior documented by Lerbs (2014) and an indirect channel operating through house prices (Melecky and Paksi, 2024). However, the puzzling role of building costs in the empirical literature persists across different dependent variables, specification designs, and sample periods.

Dynamic multipliers reveal a consistent hierarchy of adjustment speeds: income and house price effects materialize within 2–4 months, interest rate effects require 4–8 months, and unemployment and rent effects display the longest lags (8–13 months). Even the slowest channels thus operate well within business-cycle frequencies at the permit stage. Since building permits precede actual construction completions, however, the total lag from a macroeconomic impulse to the delivery of housing units can extend to several years—consistent with the long adjustment cycles traditionally attributed to residential construction.

Our findings carry several implications for housing policy design. A central lesson is that policy effectiveness is strongly conditioned by existing institutional frameworks, implying that uniform European housing policy is unlikely to be effective across countries.

Rather, the regime distinction implies country-specific policy priorities. In Germany, where interest rate transmission is strongest due to limited mortgage subsidies and developers respond to rental income signals rather than house price momentum, the most effective policy mix combines subsidized financing programs with rental market stabilization. In France, where institutional investment incentives (Pinel Law, DALO Act) drive speculative dynamics through policy design rather than financial depth, recalibrating incentive structures may be more appropriate than macroprudential tightening. In the Netherlands, where credit amplification dominates, macroprudential tools targeting credit conditions, such as loan-to-value and loan-to-income limits, are likely most effective (Kelly et al., 2019).

Austria's experience provides empirical precedent for an alternative pathway: its institutional design, combining subsidized interest rates, public low-rent housing provision, and regulatory settings dampening speculative investment, has stabilized construction independently of monetary policy conditions (Joebges et al., 2015; Luckert, 2026). Expanding non-profit housing construction on similar principles may offer a viable model for countries where market-based construction proves volatile or unresponsive to conventional policy levers.

Some limitations of our analysis suggest avenues for future research. First, while our two-regime typology provides a useful organizing framework, regime boundaries are not uniformly sharp across all transmission channels, and the classfication should be regarded as sample-dependent and potentially subject to change as institutional frameworks evolve.

Another concern is the potential for omitted variable bias. Population growth and housing preference, e.g. square meters occupied by single-person households in cities, simultaneously affect aggregate housing demand, house prices, rents, and construction activity. Omitted variables may constitute confounders that create indirect pathways biasing our coefficient estimates. Differences in population growth among countries in our sample may attribute to institutional regime differences partly reflecting differences in underlying demographic pressure. While we include real disposable income per capita as a demand-side proxy, this controls for the income dimension of demographic change but not for the quantity effect of a growing number of households seeking (urban) housing units.

Moreover, the transmission of demographic shocks to construction may not be linear as it depends critically on vacancy rates and capacity utilization in the construction sector. In high-vacancy environment such as Germany in the early 2000s, rising demand may first reduce vacancies with limited effects on rents, house prices, or new construction. In low-vacancy, capacity-constrained markets, the same demographic impulse is more likely to translate into price and rent increases, with construction responding only with considerable delay. These nonlinearities suggest that vacancy rates and capacity utilization are not merely omitted control variables but drivers of regime change we should study. As usual, however, there is a trade-off between enriching the specifications and data availability.

Finally, our in-sample estimates could be complemented by out-of-sample forecasting exercises. After several years of decline, some countries, notably Germany, are now seeing an upturn in building permits and residential construction orders. However, our empirical findings counsel caution regarding the pace of recovery. For Germany, the two channels our analysis identifies as most important, namely financing costs and construction costs, both point toward a subdued outlook. Mortgage rates are more closely linked to sovereign bond yields, which have been rising due to the 2025 fiscal stimulus package, irrespective of ECB's policy rate changes. Meanwhile, the structurally elevated building cost *level* continues to weigh on construction activity, while costs may even increase because of the war in Iran. Taken together, these factors suggest that the current recovery in German construction activity could remain rather flat.

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Tables

		Permits $\Delta^{12} \ln p$	**Income** $\Delta^{12} \ln y$	**Price** $\Delta^{12} \ln h$	**Costs** $\Delta^{12} \ln c$	**Rents** $\Delta^{12} \ln r$	**Interest** $\Delta^{12} i$	**Unemp. rate** $\Delta^{12} u$
GER	μ	−2.22	0.82	0.51	0.99	−0.75	−0.13	−0.20
	σ	16.73	1.00	5.09	2.03	1.78	0.68	0.68
FRA	μ	−0.94	1.01	2.50	1.02	−0.37	−0.10	−0.07
	σ	19.07	1.32	5.47	2.04	1.72	0.59	0.63
NED	μ	−1.27	0.64	1.36	0.53	0.10	−0.09	0.00
	σ	31.28	2.03	5.34	5.37	2.67	0.56	0.90
BEG	μ	0.41	0.54	2.02	0.37	−0.24	−0.12	−0.06
	σ	21.37	1.55	3.12	1.97	2.15	0.63	0.78
AUT	μ	−0.21	0.43	1.97	1.01	0.96	−0.08	0.05
	σ	17.29	2.03	4.38	1.66	2.84	0.76	0.86

Table 3: Descriptive Statistics: Mean and Standard Deviation of Year-over-Year Changes

Notes: μ denotes the mean and σ the standard deviation. All variables except interest rate and unemployment rate are in logarithmic differences. Year-over-year changes are calculated for the period January 2000 to June 2024. Sample size: $T = 294$ monthly observations per country.

	Income $\Delta^{12} \ln y$	**Price** $\Delta^{12} \ln h$	**Costs** $\Delta^{12} \ln c$	**Rents** $\Delta^{12} \ln r$	**Interest** $\Delta^{12} i$	**Unemp. rate** $\Delta^{12} u$
GER	0.12	0.60	−0.13	0.32	−0.40	0.01
FRA	0.32	0.50	0.12	0.14	−0.38	−0.04
NED	0.03	0.29	0.25	0.28	−0.11	−0.15
BEG	−0.01	0.22	−0.03	0.05	−0.11	0.04
AUT	−0.08	0.36	−0.15	0.12	−0.36	0.04

Notes: Contemporaneous correlation coefficients between year-over-year change in building permits ($\Delta^{12} \ln p$) and year-over-year changes in each variable. All variables except interest rates and unemployment rate are in logarithmic differences. Sample period: January 2000 to June 2024.

Table 4: Contemporaneous Correlations Between Building Permit Growth and Key Macroeconomic Variables

	GER	**FRA**	**NED**	**BEG**	**AUT**
$\log p^{ct}$	0.856	0.133	0.299	0.027	0.969
$\Delta \log p^{c}$	0.580	0.008	0.028	0.000	0.000
$\log y^{ct}$	0.598	0.832	0.656	0.420	0.789
$\Delta \log y^{c}$	0.158	0.002	0.000	0.361	0.254
$\log h^{cqt}$	0.985	0.812	0.749	0.784	0.996
$\Delta \log h^{ct}$	0.113	0.038	0.176	0.114	0.051
$\Delta^2 \log h^{c}$	0.024	0.012	0.000	0.006	0.000
$\log c^{cqt}$	0.161	0.783	0.969	0.346	0.012
$\Delta \log c^{ct}$	0.025	0.166	0.025	0.439	0.743
$\Delta^2 \log c^{c}$	0.170	0.309	0.308	0.674	0.644
$\log r^{ct}$	0.014	0.181	0.113	0.000	0.011
$\Delta \log r^{c}$	0.011	0.368	0.099	0.565	0.123
u^{ct}	0.519	0.779	0.857	0.517	0.127
Δu^{c}	0.076	0.330	0.002	0.136	0.000
i^{ct}	0.922	0.442	0.596	0.312	0.226
Δi^{c}	0.040	0.000	0.112	0.473	0.038

Notes: Bootstrap p-values for the ADF-GLS test with null hypothesis of a unit root. Superscripts indicate the deterministic specification: ^{c} = constant only; ^{ct} = constant and linear trend; ^{cqt} = constant, linear trend, and quadratic trend. The optimal lag length is determined by AIC with maximum lag length $p^{\max} = 12$. Sample period: January 2000 to June 2024.

Table 5: ADF-GLS Unit Root Test Results for Core Variables

	GER	**FRA**	**NED**	**BEG**	**AUT**
$\log p^{ct}$	Yes	Yes	Yes	Yes	Yes
$\Delta \log p^{c}$	No	No	No	No	No
$\log y^{ct}$	Yes	Yes	Yes	Yes	Yes
$\Delta \log y^{c}$	No	No	No	No	No
$\log h^{ct}$	Yes	Yes	Yes	Yes	Yes
$\Delta \log h^{ct}$	Yes	Yes	Yes	No	Yes
$\Delta^2 \log h^{c}$	No	No	No	No	No
$\log c^{ct}$	Yes	Yes	Yes	Yes	Yes
$\Delta \log c^{ct}$	No	No	Yes	No	No
$\Delta^2 \log c^{c}$	No	No	No	No	No
$\log r^{ct}$	Yes	Yes	Yes	Yes	Yes
$\Delta \log r^{c}$	No	Yes	No	No	No
u^{ct}	Yes	Yes	Yes	Yes	Yes
Δu^{c}	No	No	No	No	No
i^{ct}	Yes	Yes	Yes	Yes	Yes
Δi^{c}	Yes	No	No	No	No

Notes: KPSS test results for the null hypothesis of stationarity (level-stationary or trend-stationary). "Yes" indicates rejection of the null at the 5% level (evidence of unit root); "No" indicates failure to reject the null (stationary). Superscripts indicate the deterministic specification: ^{c} = constant only; ^{ct} = constant and linear trend. The KPSS test does not support quadratic trend specifications. Lag length is set to $p = 6$, and the bandwidth for Bartlett kernel smoothing is automatically determined. Sample period: January 2000 to June 2024 ($T = 294$ observations). Country codes: GER=Germany, FRA=France, NED=Netherlands, BEG=Belgium, AUT=Austria.

Table 6: KPSS Stationarity Test Results for Core Variables

	Model 1 ($\ln y$)	**Model 2** ($\ln y, i$)	**Model 3** ($+\Delta \ln c$)	**Model 4** ($+u$)	**Model 5** ($+\Delta \ln h$)	**Model 6** ($i, \Delta \ln c, u, \Delta \ln h, \ln r$)
GER	0.628	**0.085**	**0.032**	**0.003**	**0.002**	**0.001**
FRA	0.637	**0.000**	**0.000**	**0.000**	**0.000**	**0.000**
NED	**0.002**	**0.008**	**0.004**	**0.002**	**0.001**	**0.085**
BEG	**0.000**	**0.000**	**0.000**	**0.000**	**0.000**	**0.000**
AUT	0.954	0.557	0.498	0.239	0.396	0.370

Notes: Bootstrap p-values of the PSS F-test for the null hypothesis of no cointegration in ARDL models. The optimal lag length is selected by BIC (maximum lag length: 10). All models include a constant and trend (conditional ECM). Bootstrap procedure: wild bootstrap with 1999 replications. Model specifications are nested and cumulative: Model 1 includes income only; Model 2 adds interest rate; Model 3 adds construction cost growth; Model 4 adds unemployment; Model 5 adds house price growth; Model 6 replaces income with rents. Values in **bold** indicate rejection of the null at 10% level.

Table 7: PSS Cointegration Test Results: Bootstrap p-values for Building Permits

Country	Interest	Unemployment	House Price	Income	Costs	Rents
Germany	-0.20^{***} to -0.26^{***} (Models 2-5)	-0.08^{***} to -0.09^{***} (Models 4-6)	Insignificant (Models 5-6)	5.42 to 5.45 (Models 4-5, insig.)	Negative but insig. -44 to -16 (Models 3-6)	6.14^{**} (Model 6)
France	-0.14^{***} to -0.23^{***} (Models 2-5)	0.05^{**} (Model 5) Insig. (Models 4, 6)	15.5^{**} to 33.3^{***} (Models 5-6)	4.87^{***} to 5.26^{***} (Models 2-5)	Positive sig. 11.3^{**} to 13.8^{**} (Models 3-5) M6 borderline	2.72^{***} (Model 6)
Netherlands	-0.12^{*} (Model 2, 10% sig.) Insig. (Models 3-6) M5 positive	Insignificant (Models 4-5) M6 borderline	31.0^{***} to 60.9^{***} (Models 5-6)	8.26^{***} to 10.6^{***} (Models 1-5)	Positive sig. 12.1^{**} to 18.3^{***} (Models 3-5) M6 insig.	-5.32^{**} (Model 6)
Belgium	-0.09^{***} (M4) -0.05^{*} (M6) Other models insig.	Insignificant (Models 4-6)	18.7^{*} (Model 5, 10% sig.) M6 insig.	Insignificant (Models 1-5)	Insignificant (Models 3-6)	Insignificant (Model 6)
Austria	*Excluded due to lack of cointegration (see Section 4.2)*					

Notes: $^{*}p < 0.10,^{**} p < 0.05,^{***} p < 0.01$ based on bootstrap t-statistics (1999 replications).
Values represent long-run coefficients. For i_t and u_t, values are semi-elasticities (effect of 1 pp change).
For Costs and House Prices, inputs are $\Delta \ln c$ and $\Delta \ln h$; coefficients represent semi-elasticities w.r.t. monthly growth rates.
"Insig." = insignificant; "M" = Model number; "borderline" = t-stat close to critical value but not reaching 10% level.

Table 8: Long-Run Multipliers on Building Permits Across European Countries Using The ARDL Model

Country	**Interest Rate**	**Unemployment**	**House Price**
Germany	-0.15 after 16m (sig.), Models 3-5, from 4-6m Model 6: n.s.	Negative sig. after 11-13m Models 4-6 Initial positive (counter-int.) lasts few months	No significant effects Models 5-6
France	-0.175 after 16m (sig.), Models 3-4, from 4m Model 5: from 8m Model 6: n.s.	Positive (counter-int.) High uncertainty M5 positive after 4m M4 borderline	$\sim$10.0 at 4m, $\sim$28.0 at 16m Model 6 (sig. from 3-4m) Model 5 similar
Netherlands	Model 3: borderline from 10m Models 4-6: n.s. M5 positive (implausible)	Immediate neg. -0.10 (1m) M6 lasting minor neg. (sig.) M5 pos. after 7m (almost sig. at 12m)	Large multipliers (sig.) Larger than France From 3-4m onwards
Belgium	Models 4, 6: neg. after 8m Mean -0.1 (sig.) Other models similar but n.s.	M6 borderline neg. after 3m Other models n.s.	$\sim$10.0 at 4m, $\sim$20.0 at 16m Model 5 only (sig.) From 3m onwards
Austria†	M4: sig. neg. from 3m M3: sig. from 7m M5-6: n.s.	M4: neg. -0.03 (sig.) from 4m M5: borderline at 5m M6: n.s.	1.0 to 2.0 (sig.) Models 5-6, immediate Lasting throughout 16m (brief exceptions M6, m8-13)

Notes: sig. = significant at 10% level; n.s. = not significant; m = months. Values represent mean multipliers after 16 months unless otherwise noted. † Austria estimated in first differences; multipliers represent effects on permit *growth rates*.

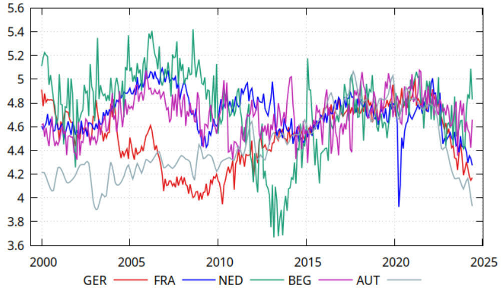
Table 9: Dynamic Multipliers on Building Permits Across European Countries (Interest Rate, Unemployment, and House Price)

Table 10: Dynamic Multipliers on Building Permits Across European Countries (Income, Costs, and Rents)

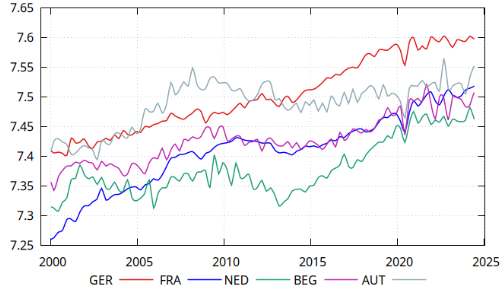
Country	**Income**	**Costs**	**Rents**
Germany	2.0 at 4m, 4.4 at 16m Models 4-5 (sig. from 4m) Just at margin of sig. Model 3: n.s.	Negative, all models Large uncertainty Just at border of sig. M6: n.s.	$\sim$2.0 at 8m, $\sim$5.0 at 16m Sig. from 8m High initial uncertainty
France	Immediate $\sim$6.0, correction Then $\sim$3.0 at 4m, $\sim$5.0 at 16m Models 3-5 (sig. from 3m)	Positive (counter-int.) Models 3-6 (sig.) From 1m onwards	$\sim$1.5 (sig.) From 8m onwards Large initial uncertainty
Netherlands	Initial positive, correction $\sim$4.0 at 2m, $\sim$9.0 at 16m All models (sig. from 2m)	Tendency positive M6: n.s. Others: borderline	Never sig. for 2 consecutive m Negative after 10m but insignificant
Belgium	M4: pos. after 9m ($\sim$2.0) Other models: high uncertainty, n.s.	M3-5: negative mean (n.s.) M6: positive mean (n.s.) Expected sign only M3-5	Positive mean $\sim$1.5 after 4m Borderline sig. throughout
Austria†	Counter-intuitive negative Borderline first 2m only No robust effects	All models negative (sig.) From 9-10m onwards Comparable multipliers across models	Brief sig. months 6-8 ($\sim$-2.0) Otherwise n.s.

Notes: sig. = significant at 10% level; n.s. = not significant; m = months. Values represent mean multipliers after 16 months unless otherwise noted. † Austria estimated in first differences; multipliers represent effects on permit *growth rates*.

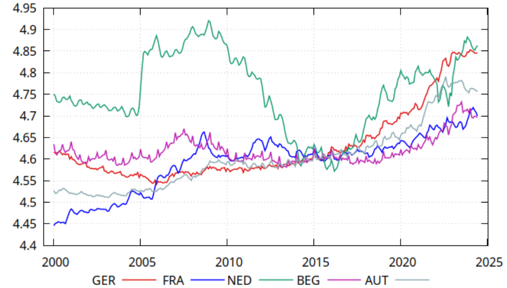
Figures



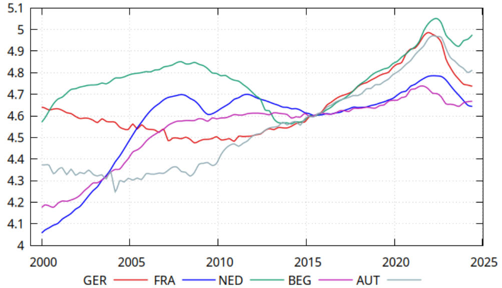
(a) Log of permits p



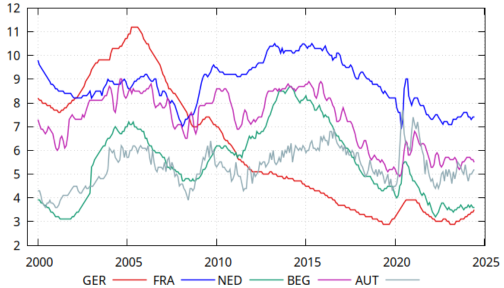
(b) Log of income y



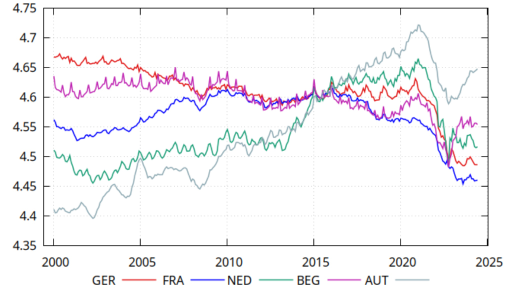
(c) Log of building costs c



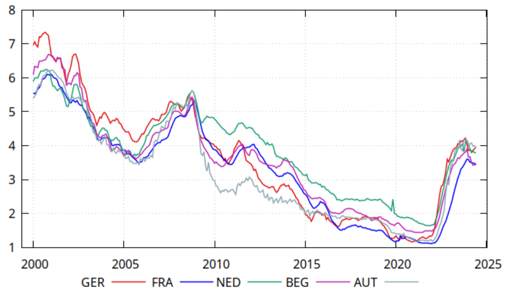
(d) Log of house prices h



(e) Unemployment rate u

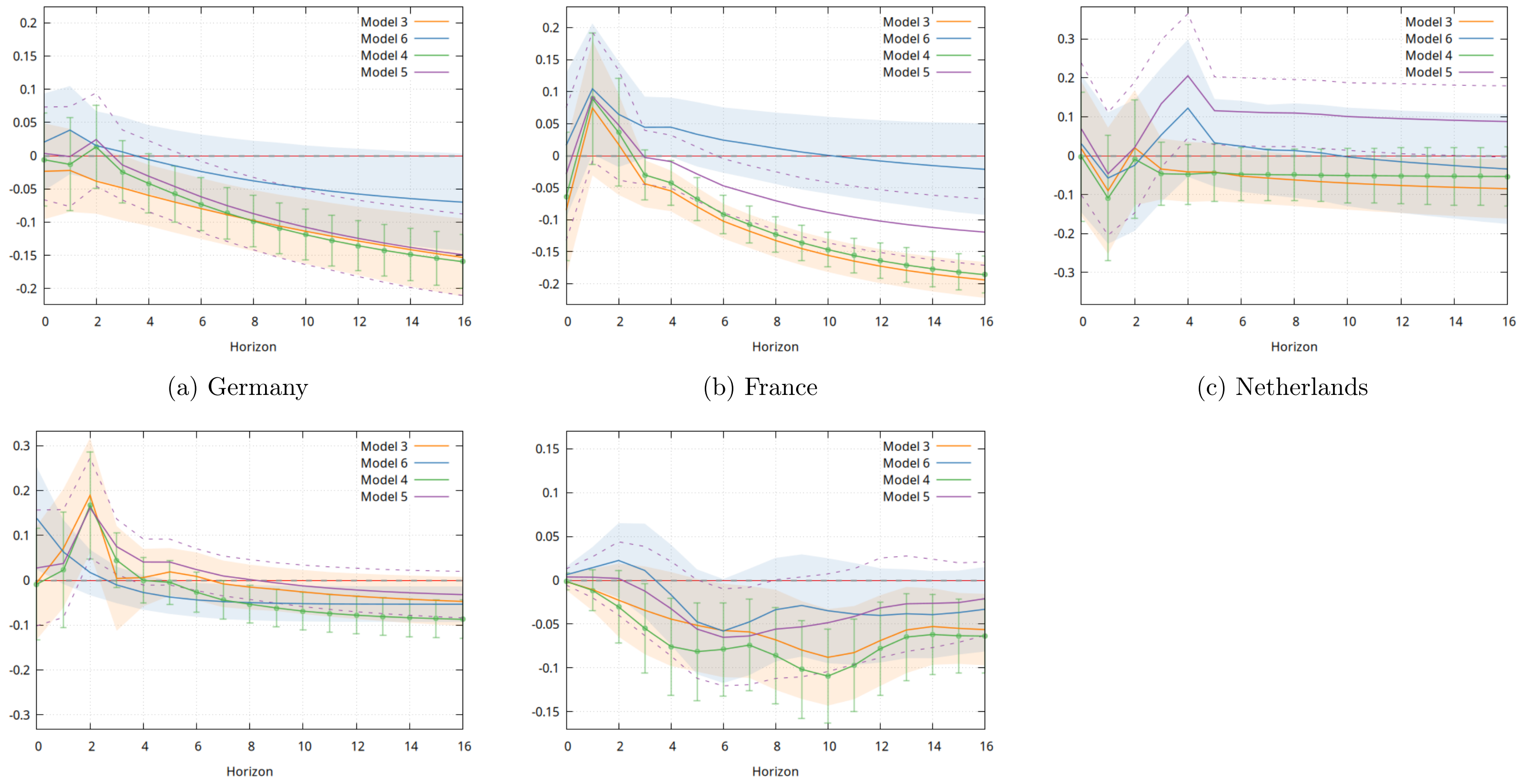


(f) Log of rent prices r



(g) Interest rate i

Figure 2: Time-series Overview

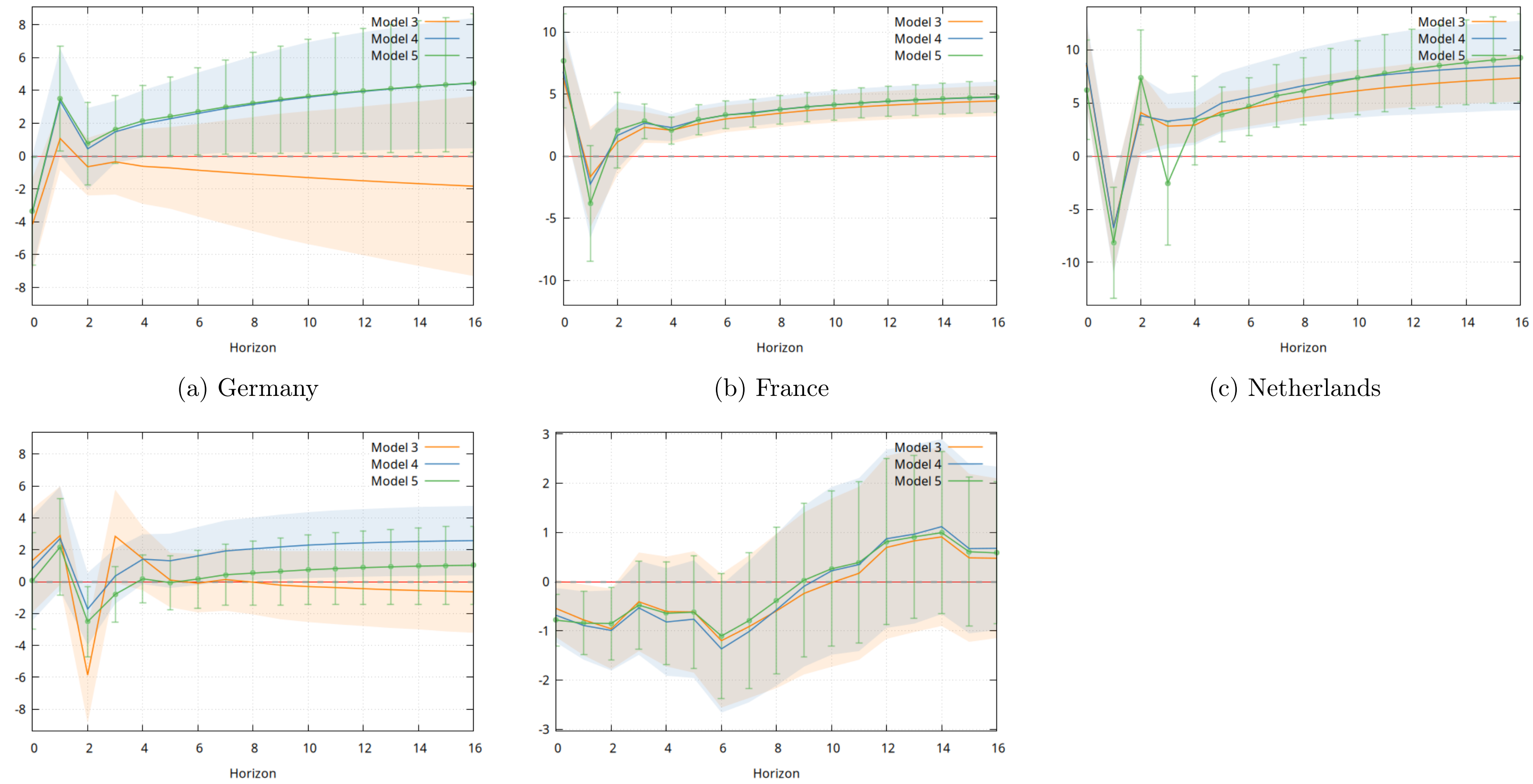


(a) Germany (b) France (c) Netherlands

(d) Belgium (e) Austria

Note: Plots show the mean and 84% confidence intervals of wild-bootstrapped dynamic multipliers using 1999 iterations. For Germany, France, the Netherlands, and Belgium, multipliers are estimated from the ARDL-ECM in levels, where cointegration is supported (see Section 4.2). For Austria, where cointegration tests fail to reject the null of no long-run relationship, we report results from a first-difference specification; these estimates capture short- to medium-run dynamics but do not have a long-run equilibrium interpretation.

Figure 3: Cumulative dynamic multipliers from the ARDL model in response to a unit interest rate increase. For Germany, France, the Netherlands, and Belgium, the response of log permits to the interest rate is shown. For Austria, the response of permit growth to a change in the interest rate is shown based on a first-difference specification.

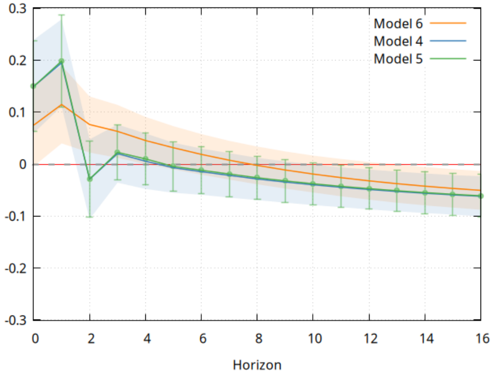


(a) Germany (b) France (c) Netherlands

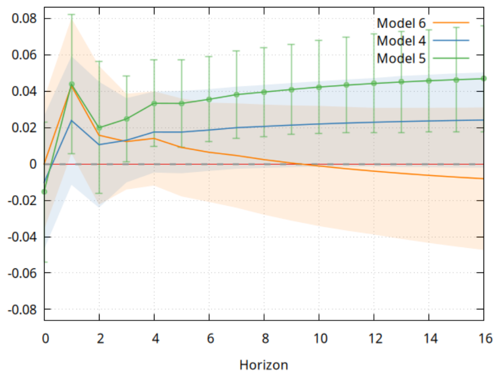
(d) Belgium (e) Austria

Note: Plots show the mean and 84% confidence intervals of wild-bootstrapped dynamic multipliers using 1999 iterations. For Germany, France, the Netherlands, and Belgium, multipliers are estimated from the ARDL-ECM in levels, where cointegration is supported (see Section 4.2). For Austria, where cointegration tests fail to reject the null of no long-run relationship, we report results from a first-difference specification; these estimates capture short- to medium-run dynamics but do not have a long-run equilibrium interpretation.

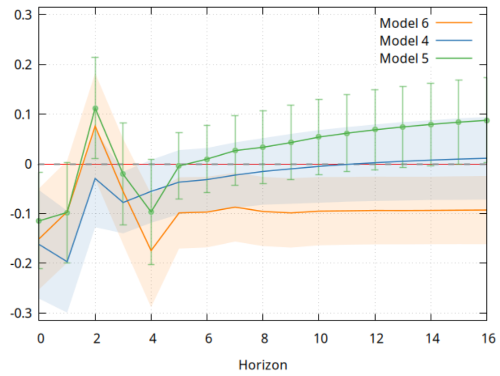
Figure 4: Cumulative dynamic multipliers from the ARDL model in response to a permanent unit increase in the logarithm of income. For Germany, France, the Netherlands, and Belgium, the response of log permits to log income is shown. For Austria, the response of permit growth to income growth is shown based on a first-difference specification.



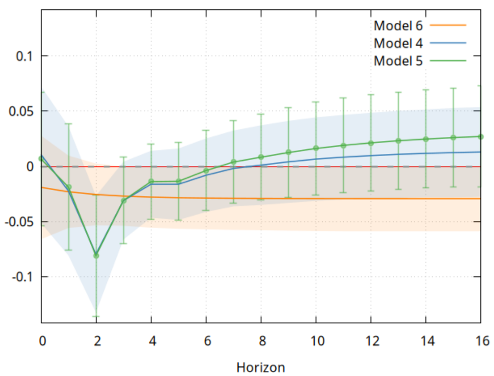
(a) Germany



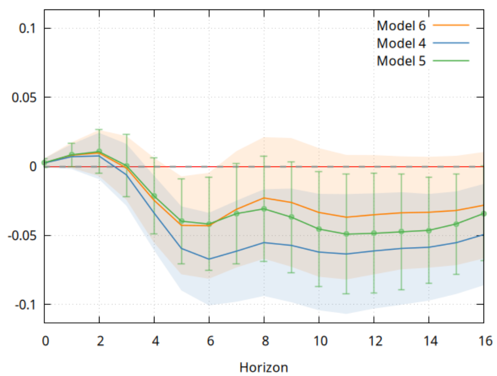
(b) France



(c) Netherlands



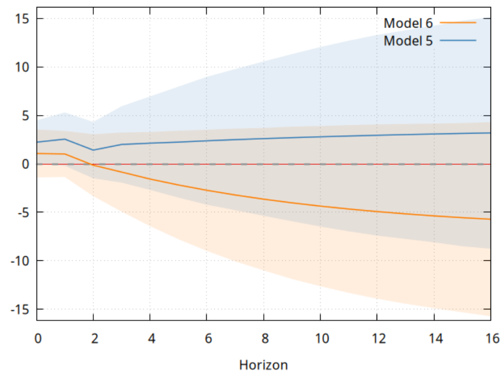
(d) Belgium



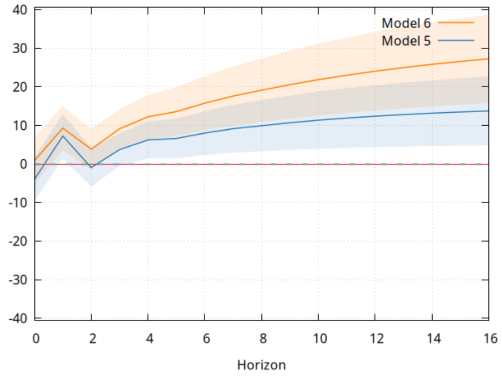
(e) Austria

Note: Plots show the mean and 84% confidence intervals of wild-bootstrapped dynamic multipliers using 1999 iterations. For Germany, France, the Netherlands, and Belgium, multipliers are estimated from the ARDL-ECM in levels, where cointegration is supported (see Section 4.2). For Austria, where cointegration tests fail to reject the null of no long-run relationship, we report results from a first-difference specification; these estimates capture short- to medium-run dynamics but do not have a long-run equilibrium interpretation.

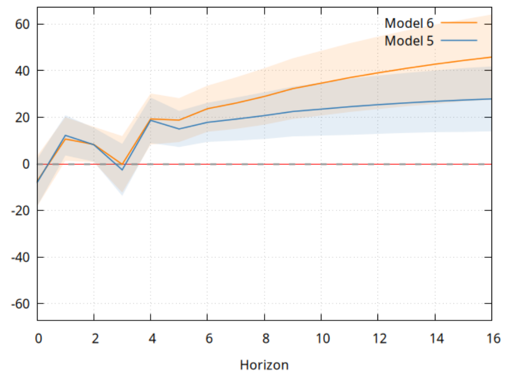
Figure 5: Cumulative dynamic multipliers from the ARDL model in response to a permanent unit unemployment rate increase . For Germany, France, the Netherlands, and Belgium, the response of log permits to unemployment rate is shown. For Austria, the response of permit growth to a month-to-month change in the unemployment rate is shown based on a first-difference specification.



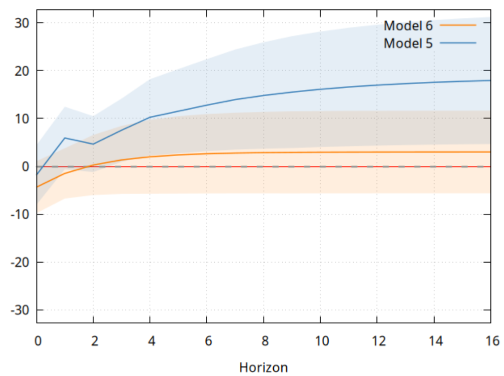
(a) Germany



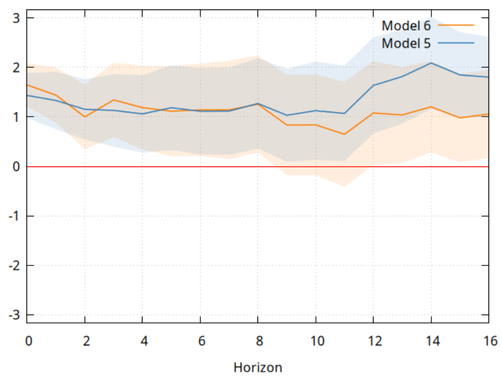
(b) France



(c) Netherlands



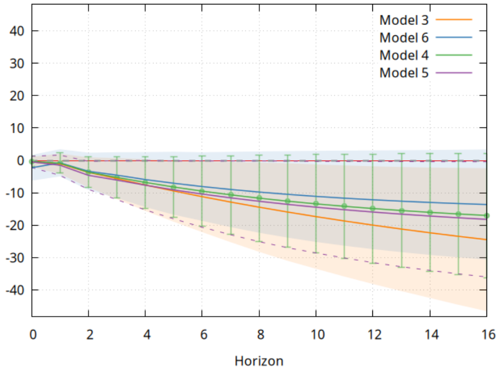
(d) Belgium



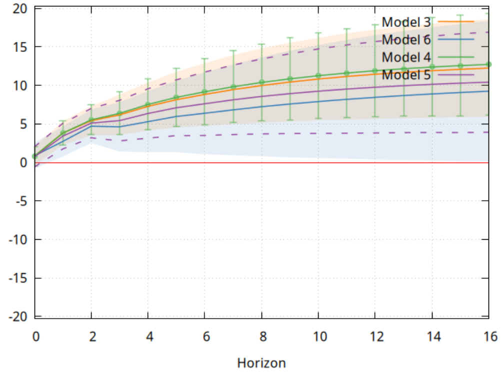
(e) Austria

Note: Plots show the mean and 84% confidence intervals of wild-bootstrapped dynamic multipliers using 1999 iterations. For Germany, France, the Netherlands, and Belgium, multipliers are estimated from the ARDL-ECM in levels, where cointegration is supported (see Section 4.2). For Austria, where cointegration tests fail to reject the null of no long-run relationship, we report results from a first-difference specification; these estimates capture short- to medium-run dynamics but do not have a long-run equilibrium interpretation.

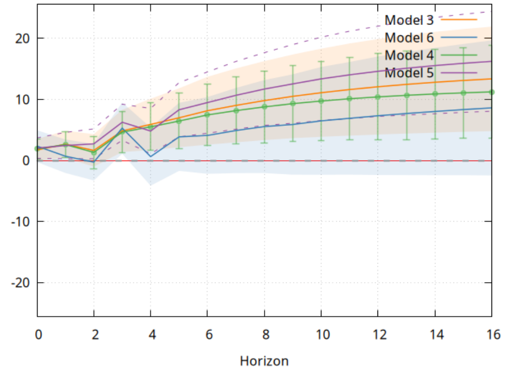
Figure 6: Cumulative dynamic multipliers from the ARDL model in response to a permanent unit increase in house prices. For Germany, France, the Netherlands, and Belgium, the response of log permits to log house prices is shown. For Austria, the response of permit growth to growth in house prices is shown based on a first-difference specification.



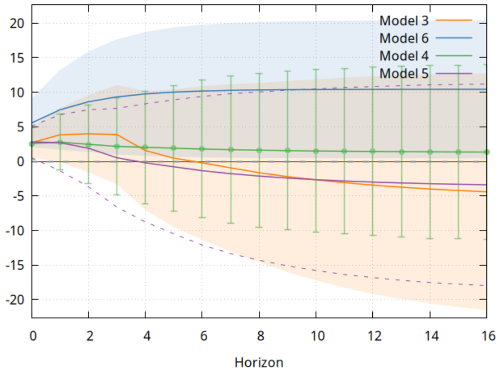
(a) Germany



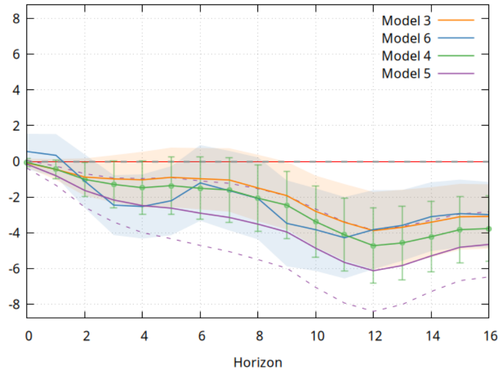
(b) France



(c) Netherlands



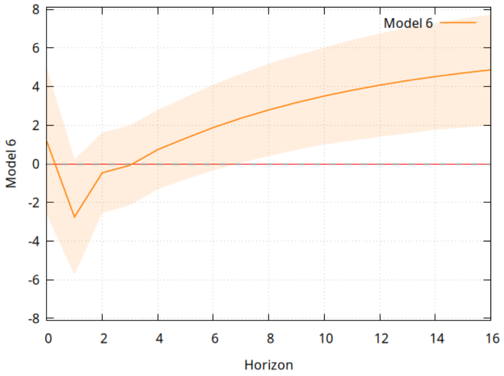
(d) Belgium



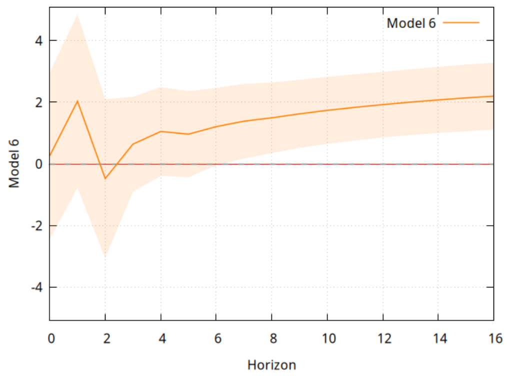
(e) Austria

Note: Plots show the mean and 84% confidence intervals of wild-bootstrapped dynamic multipliers using 1999 iterations. For Germany, France, the Netherlands, and Belgium, multipliers are estimated from the ARDL-ECM in levels, where cointegration is supported (see Section 4.2). For Austria, where cointegration tests fail to reject the null of no long-run relationship, we report results from a first-difference specification; these estimates capture short- to medium-run dynamics but do not have a long-run equilibrium interpretation.

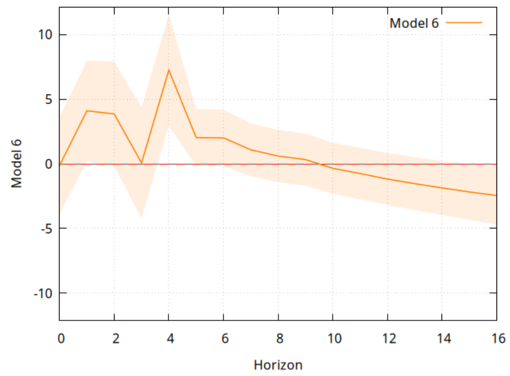
Figure 7: Cumulative dynamic multipliers from the ARDL model in response to a permanent unit increase in building costs. For Germany, France, the Netherlands, and Belgium, the response of log permits to log building costs is shown. For Austria, the response of permit growth to growth in building costs is shown based on a first-difference specification.



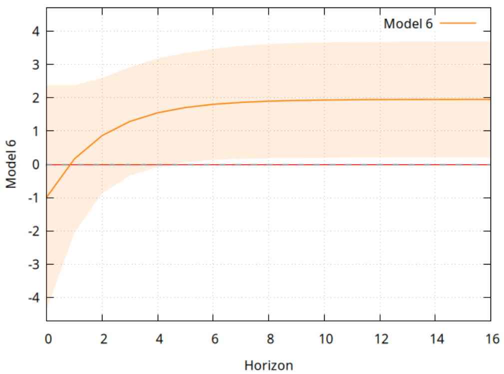
(a) Germany



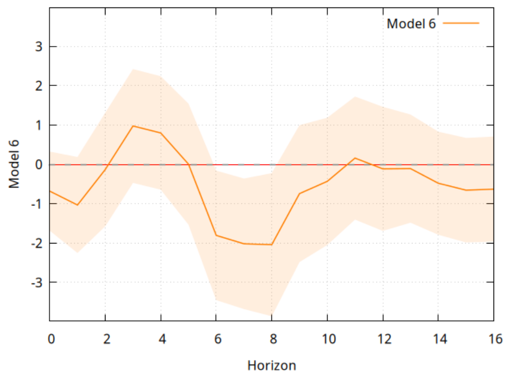
(b) France



(c) Netherlands



(d) Belgium



(e) Austria

Note: Plots show the mean and 84% confidence intervals of wild-bootstrapped dynamic multipliers using 1999 iterations. For Germany, France, the Netherlands, and Belgium, multipliers are estimated from the ARDL-ECM in levels, where cointegration is supported (see Section 4.2). For Austria, where cointegration tests fail to reject the null of no long-run relationship, we report results from a first-difference specification; these estimates capture short- to medium-run dynamics but do not have a long-run equilibrium interpretation.

Figure 8: Cumulative dynamic multipliers from the ARDL model in response to a permanent unit increase in rent prices. For Germany, France, the Netherlands, and Belgium, the response of log permits to log rents is shown. For Austria, the response of permit growth to rent growth is shown based on a first-difference specification.

Summary of Correlation Analysis of Building Permits

The following bullet points summarize the main results of the contemporaneous correlations between year-over-year growth in building permits, and key covariates as reported in Table 4.

- House prices: Permits co-move positively with house prices in all five countries: GER (0.60), FRA (0.50), AUT (0.36), NED (0.29), BEG (0.22). The strongest contemporaneous association appears in GER and FRA.
- Interest rates: Correlations are uniformly negative and sizable in GER (-0.40), FRA (-0.38), AUT (-0.36), and weaker in NED/BEG (-0.11 each), consistent with the financing-cost channel.
- Rents: Permits show small positive co-movement with rents in GER (0.32), NED (0.28), AUT (0.12), FRA (0.14), and are near zero in BEG (0.05), suggesting heterogeneous rental-market links.
- Income: Contemporaneous links are modest: FRA (0.32) is the only clear positive association; GER (0.12) is small; NED (0.03), BEG (-0.01), AUT (-0.08) are negligible, in line with income being a slower-moving determinant.
- Construction costs: Mixed, small and positive correlations in NED (0.25) and FRA (0.12); small and negative in GER (-0.13) and AUT (-0.15), BEG (-0.03) negative.
- Unemployment: Near-zero contemporaneous correlations in GER (0.01), FRA (-0.04), BEG (0.04), AUT (0.04), and small in NED (-0.15), indicating weak immediate co-movement.

Appendix: Data and Tables

Tables

	Model 1 ($\ln y$)	**Model 2** ($\ln y, i$)	**Model 3** ($+\Delta \ln c$)	**Model 4** ($+u$)	**Model 5** ($+\Delta \ln h$)	**Model 6** ($i, \Delta \ln c, u, \Delta \ln h, \ln r$)
GER	0.509	**0.071**	**0.086**	**0.005**	**0.004**	**0.002**
FRA	0.549	**0.000**	**0.000**	**0.000**	**0.000**	**0.001**
NED	**0.003**	**0.001**	**0.002**	**0.000**	**0.000**	**0.019**
BEG	**0.000**	**0.000**	**0.000**	**0.000**	**0.000**	**0.000**
AUT	0.914	0.383	0.580	0.714	0.654	0.631

Notes: Bootstrap p-values of the BDM t-test for the null hypothesis of no cointegration in ARDL models. The test is based on the t-statistic of the error-correction coefficient ($\rho = 0$ under the null). The optimal lag length is selected by BIC (maximum lag length: 10). All models include a constant and trend (conditional ECM). Bootstrap procedure: wild bootstrap with 1999 replications. Model specifications are nested and cumulative: Model 1 includes income only; Model 2 adds interest rate; Model 3 adds construction cost growth; Model 4 adds unemployment; Model 5 adds house price growth; Model 6 replaces income with rents. Country codes: GER=Germany, FRA=France, NED=Netherlands, BEG=Belgium, AUT=Austria. **Values in bold indicate rejection of the null at 10% level ($p \leq 0.10$).**

Table A1: BDM Cointegration Test Results: Bootstrap p-values for Building Permits

	Model 1 ($\ln y$)	**Model 2** ($+i$)	**Model 3** ($+\Delta \ln c$)	**Model 4** ($+u$)	**Model 5** ($+\Delta \ln h$)	**Model 6** ($i, \Delta \ln c, u, \Delta \ln h, \ln r$)
Error correct. (ρ)	-0.044	-0.091	-0.066	-0.142	-0.140	-0.146
Income ($\ln y$)	1.258	-1.891	-3.230	5.445	5.418	—
t-stat	(0.126)	(-0.310)	(-0.493)	(1.522)	(1.437)	—
Interest rate (i)	—	-0.217^{***}	-0.259^{***}	-0.210^{***}	-0.202^{***}	-0.090
t-stat	—	(-3.001)	(-3.574)	(-4.810)	(-3.633)	(-1.483)
Costs ($\Delta \ln c$)	—	—	-43.526	-21.778	-23.510	-15.986
t-stat	—	—	(-1.618)	(-1.156)	(-1.243)	(-1.143)
Unemployment (u)	—	—	—	-0.088^{***}	-0.090^{***}	-0.080^{***}
t-stat	—	—	—	(-2.755)	(-2.649)	(-2.607)
House prices ($\Delta \ln h$)	—	—	—	—	3.844	-6.989
t-stat	—	—	—	—	(0.347)	(-0.847)
Rents ($\ln r$)	—	—	—	—	—	6.136^{**}
t-stat	—	—	—	—	—	(2.587)
p	2	2	2	3	3	2
SC[1] (p-value)	0.465	0.817	0.829	0.269	0.398	0.694
SC[12] (p-value)	0.150	0.194	0.261	0.270	0.362	0.839
White (p-value)	0.240	0.139	0.163	0.595	0.682	0.288
Adjusted R^2	0.921	0.924	0.924	0.928	0.928	0.927

Notes: Bootstrap long-run coefficient estimates with t-statistics in parentheses. ρ denotes the bootstrap mean error-correction coefficient (speed of adjustment toward long-run equilibrium). Significance levels: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$. The optimal lag length is selected by AIC (maximum lag: 10). All models include a constant and trend (conditional ECM). Bootstrap procedure: wild bootstrap with 1999 replications. Diagnostics: p = optimal lag order; SC[1] and SC[12] = Breusch-Godfrey serial correlation test p-values for orders 1 and 12; White = White heteroscedasticity test p-value. Target variable: log building permits.

Table A2: Long-Run Coefficient Estimates: Germany (GER)

	Model 1 ($\ln y$)	**Model 2** ($+i$)	**Model 3** ($+\Delta \ln c$)	**Model 4** ($+u$)	**Model 5** ($+\Delta \ln h$)	**Model 6** ($i, \Delta \ln c, u, \Delta \ln h, \ln r$)
Error correct. (ρ)	-0.040	-0.216	-0.236	-0.240	-0.234	-0.166
Income ($\ln y$)	-1.032	4.868^{***}	4.865^{***}	5.161^{***}	5.255^{***}	—
t-stat	(-0.165)	(4.607)	(5.130)	(5.476)	(5.240)	—
Interest rate (i)	—	-0.229^{***}	-0.221^{***}	-0.212^{***}	-0.141^{***}	-0.046
t-stat	—	(-8.252)	(-8.686)	(-8.384)	(-3.601)	(-0.761)
Costs ($\Delta \ln c$)	—	—	13.283^{**}	13.767^{***}	11.291^{**}	10.754
t-stat	—	—	(2.574)	(2.630)	(2.117)	(1.344)
Unemployment (u)	—	—	—	0.026	0.050^{**}	-0.016
t-stat	—	—	—	(1.268)	(2.172)	(-0.480)
House prices ($\Delta \ln h$)	—	—	—	—	15.541^{**}	33.290^{***}
t-stat	—	—	—	—	(2.042)	(3.279)
Rents ($\ln r$)	—	—	—	—	—	2.720^{***}
t-stat	—	—	—	—	—	(3.104)
p	3	3	3	3	3	3
SC[1] (p-value)	0.058	0.052	0.061	0.025	0.026	0.029
SC[12] (p-value)	0.146	0.328	0.322	0.346	0.172	0.342
White (p-value)	0.504	0.289	0.125	0.380	0.782	0.574
Adjusted R^2	0.890	0.904	0.909	0.908	0.910	0.906

Notes: Bootstrap long-run coefficient estimates with t-statistics in parentheses. ρ denotes the bootstrap mean error-correction coefficient (speed of adjustment toward long-run equilibrium). Significance levels: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$. The optimal lag length is selected by AIC (maximum lag: 10). All models include a constant and trend (conditional ECM). Bootstrap procedure: wild bootstrap with 1999 replications. Diagnostics: p = optimal lag order; SC[1] and SC[12] = Breusch-Godfrey serial correlation test p-values for orders 1 and 12; White = White heteroscedasticity test p-value. Target variable: log building permits.

Table A3: Long-Run Coefficient Estimates: France (FRA)

	Model 1 ($\ln y$)	**Model 2** ($+i$)	**Model 3** ($+\Delta \ln c$)	**Model 4** ($+u$)	**Model 5** ($+\Delta \ln h$)	**Model 6** ($i, \Delta \ln c, u, \Delta \ln h, \ln r$)
Error correct. (ρ)	-0.177	-0.236	-0.217	-0.278	-0.272	-0.190
Income ($\ln y$)	9.076***	8.864***	8.258***	9.214***	10.613***	—
t-stat	(4.427)	(4.846)	(4.756)	(2.705)	(3.184)	—
Interest rate (i)	—	−0.116*	−0.096	−0.055	0.078	−0.076
t-stat	—	(−1.724)	(−1.540)	(−0.915)	(1.091)	(−0.627)
Costs ($\Delta \ln c$)	—	—	15.092**	12.126**	18.285***	11.485
t-stat	—	—	(2.214)	(2.061)	(2.820)	(1.067)
Unemployment (u)	—	—	—	0.021	0.112	−0.090
t-stat	—	—	—	(0.320)	(1.599)	(−1.622)
House prices ($\Delta \ln h$)	—	—	—	—	31.010***	60.863***
t-stat	—	—	—	—	(2.741)	(3.502)
Rents ($\ln r$)	—	—	—	—	—	−5.323**
t-stat	—	—	—	—	—	(−2.335)
p	3	3	3	3	5	5
SC[1] (p-value)	0.470	0.413	0.487	0.630	0.605	0.277
SC[12] (p-value)	0.608	0.749	0.922	0.840	0.727	0.656
White (p-value)	0.081	0.158	0.099	0.064	0.325	0.882
Adjusted R^2	0.770	0.770	0.774	0.779	0.791	0.778

Notes: Bootstrap long-run coefficient estimates with t-statistics in parentheses. ρ denotes the bootstrap mean error-correction coefficient (speed of adjustment toward long-run equilibrium). Significance levels: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$. The optimal lag length is selected by AIC (maximum lag: 10). All models include a constant and trend (conditional ECM). Bootstrap procedure: wild bootstrap with 1999 replications. Diagnostics: p = optimal lag order; SC[1] and SC[12] = Breusch-Godfrey serial correlation test p-values for orders 1 and 12; White = White heteroscedasticity test p-value. Target variable: log building permits.

Table A4: Long-Run Coefficient Estimates: Netherlands (NED)

	Model 1 ($\ln y$)	**Model 2** ($+i$)	**Model 3** ($+\Delta \ln c$)	**Model 4** ($+u$)	**Model 5** ($+\Delta \ln h$)	**Model 6** ($i, \Delta \ln c, u, \Delta \ln h, \ln r$)
Error correct. (ρ)	-0.199	-0.264	-0.237	-0.343	-0.281	-0.392
Income ($\ln y$)	0.891	-0.766	-0.862	2.673	1.154	—
t-stat	(0.422)	(-0.419)	(-0.440)	(1.629)	(0.645)	—
Interest rate (i)	—	-0.063	-0.061	-0.093^{***}	-0.040	-0.054^{*}
t-stat	—	(-1.612)	(-1.410)	(-2.894)	(-1.089)	(-1.915)
Costs ($\Delta \ln c$)	—	—	-5.557	1.269	-3.700	10.425
t-stat	—	—	(-0.414)	(0.136)	(-0.345)	(1.555)
Unemployment (u)	—	—	—	0.015	0.032	-0.029
t-stat	—	—	—	(0.526)	(0.948)	(-1.362)
House prices ($\Delta \ln h$)	—	—	—	—	18.731^{*}	3.010
t-stat	—	—	—	—	(1.883)	(0.476)
Rents ($\ln r$)	—	—	—	—	—	1.952
t-stat	—	—	—	—	—	(1.522)
p	4	4	4	3	3	1
SC[1] (p-value)	0.299	0.288	0.281	0.758	0.579	0.031
SC[12] (p-value)	0.167	0.116	0.110	0.553	0.151	0.095
White (p-value)	0.696	0.689	0.849	0.604	0.500	0.623
Adjusted R^2	0.561	0.571	0.570	0.570	0.568	0.548

Notes: Bootstrap long-run coefficient estimates with t-statistics in parentheses. ρ denotes the bootstrap mean error-correction coefficient (speed of adjustment toward long-run equilibrium). Significance levels: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$. The optimal lag length is selected by AIC (maximum lag: 10). All models include a constant and trend (conditional ECM). Bootstrap procedure: wild bootstrap with 1999 replications. Diagnostics: p = optimal lag order; SC[1] and SC[12] = Breusch-Godfrey serial correlation test p-values for orders 1 and 12; White = White heteroscedasticity test p-value. Target variable: log building permits.

Table A5: Long-Run Coefficient Estimates: Belgium (BEG)

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