Price Setting During a Currency Changeover*

Fernando Alvarez
University of Chicago and NBER

David Argente
Yale University and NBER

Alberto Cavallo
Harvard Business School and NBER

Francesco Lippi
LUISS University and EIEF

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Abstract

We use the euro cash changeover, a large-scale and purely nominal reform that requires all firms to redenominate posted prices, to distinguish among theories of price rigidity. We develop a nonstationary menu-cost model in which firms anticipate the changeover date and may combine currency adoption with real price adjustment. The announcement makes the firm's inaction region time-varying, a singular control problem with a moving boundary for which closed-form characterizations are rarely available. We solve this problem analytically by perturbation, recovering the whole path of the boundary, the date at which firms begin adopting the new currency, and the implied pricing moments. The model predicts that, as the changeover approaches, firms adjust more frequently but by smaller amounts, compressing the distribution of price changes. We test these predictions using historical CPI micro data from Austria, Finland, Greece, and Slovakia, a UK benchmark, and daily price data from the recent changeovers in Croatia and Bulgaria. Across changeovers, adjustment frequency rises while the size and dispersion of price changes fall. The model adds a single parameter to the standard menu-cost setup, the cost of pricing in the wrong currency. Calibrating it to the observed rise in adjustment frequency, the model reproduces the decline in the size of price changes and its subsequent recovery, neither of which is targeted. The evidence supports state-dependent price-setting models over models in which nominal redenomination leaves real pricing behavior unchanged.

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1 Introduction

We study price-setting behavior around the euro cash changeover, the reform that replaced national currencies with euro banknotes and coins in twenty-one countries, in waves running from 2002 to 2026. The changeover represents a rare large-scale episode in which all price tags in an economy are replaced on a predetermined date. Because the reform is purely *nominal*, it provides a clean natural experiment for distinguishing among competing models of price rigidity. In a frictionless world, a changeover would produce only nominal changes, leaving all pricing moments, including the frequency, size, and dispersion of adjustments, completely unaffected. In contrast, when firms face a fixed cost of changing prices, the impending changeover alters optimal pricing behavior, affecting the timing and size of price changes and, as a result, relative prices.

To fix ideas, consider two leading accounts of price stickiness: menu-cost and observation-cost models. Under menu costs, changing a price and changing its denomination are the same act: a firm that resets its price also redenominates it. Switching early is costly because euro cash was not yet in circulation, so a firm quoting in euros complicates transactions for customers still paying in the legacy currency; symmetrically, retaining legacy prices after the conversion is costly once that currency ceases to be legal tender.[1] The second cost begins at the conversion date; the first is borne only until then, and so shrinks as the date nears. Adjusting therefore becomes more attractive, the range of price gaps a firm tolerates narrows, and price changes become more frequent and smaller. Under observation costs the same reform is inert: relabeling requires no information about the desired price, so the firm converts the number on the tag and leaves its schedule for gathering information—and hence its real pricing—untouched.

These are sharply different predictions about the same purely nominal event. Which friction is at work determines how much output responds to a nominal shock. When adjustment is state dependent, a large shock pushes many firms past their adjustment threshold at once, prices move quickly, and the shock is close to neutral; when the timing of adjustment does not respond to the state, the same shock passes through slowly and its real effects are large and persistent. The difficulty is that the two classes of model are nearly indistinguishable in the data usually brought to bear on them. Alvarez, Lippi and Passadore (2016b) show that state- and time-dependent models calibrated to the same steady-state moments deliver nearly identical output responses to small monetary shocks, with differences emerging only once shocks are large, and Auclert, Rigato, Rognlie and Straub (2024) prove that the first-order inflation dynamics of a broad class of menu-cost models coincide with those of a suitably

[1] We follow Hobijn, Ravenna and Tambalotti (2006), who model both as a flow cost of "currency mismatch": a per-period cost for euro pricing before the conversion, and additional cost for legacy pricing after it.

parameterized Calvo model. Precisely where the models agree, small shocks and first-order dynamics, is where most of the evidence lies. The changeover sits on the other side of that divide: it is a large and anticipated shift in the cost of adjustment, and it therefore activates the extensive margin that the equivalence results neutralize.

Motivated by this distinction, we develop a menu-cost model of a known currency changeover. Changing a price, and thereby redenominating it, requires paying a menu cost ψ; pricing in the wrong currency—euros before the conversion date T, or the legacy currency after—costs a flow ϕ, so both the mismatch and its duration enter the firm's objective, as in Hobijn, Ravenna and Tambalotti (2006).

The announcement makes the problem nonstationary: value functions, adjustment boundaries, and the distribution of price gaps all depend on the time remaining until T. The solution method is itself a contribution. Classic characterizations of Ss pricing, Barro (1972) and Dixit (1991), deliver closed forms for a *constant* inaction band in a stationary environment. A known future date breaks stationarity: the value function solves a partial differential equation whose free boundary moves with the time remaining until the changeover, and such problems are almost always handled numerically. We show that linearizing the problem in the mismatch cost ϕ maps it into a heat equation on a fixed domain with time-varying Dirichlet data, which admits a closed-form eigenfunction solution; the perturbation approach follows Alvarez, Lippi and Souganidis (2023), who introduced these methods to the analysis of price-setting problems. The result is an explicit description of the entire transition: the path of the inaction boundary, the first adoption date T_0, and the implied frequency and size of adjustments. As the discount rate vanishes, the lead time of early adoption is

$$T - T_0 \approx \frac{8 \log 3}{\pi^2 N},$$

a pure constant divided by N, the sector's steady-state frequency of price adjustment, and nothing else; the compression of the band at the changeover likewise depends on the primitives only through N and the size of the mismatch cost.

Because these theoretical objects map directly into observable moments, we test the model using several sources of micro price data. For the earlier changeovers, obtaining the underlying price quotes is difficult because the data are confidential and in many countries are no longer available. We were able to obtain confidential CPI micro data from Austria, Finland, Greece, and Slovakia. We complement these data with CPI micro data from the United Kingdom, which did not adopt the euro and serves as a benchmark. We also use high-frequency scraped prices from Croatia around its 2023 changeover and retailer-level micro data from Bulgaria's 2026 changeover, covering more than 200 retailers before and after the conversion. Together, these datasets allow us to trace the dynamics of price adjustment around the changeover and to compare them across countries and sectors. To the best of our knowledge, this is the first paper to combine confidential CPI micro data from several countries in a common panel framework for analysis using panel-data methods.

The changeovers display a common and distinctive pattern. The frequency of price adjustments increases sharply around the introduction of euro cash. The rise is not confined to one direction, as both price increases and price decreases become more frequent. At the same time, the average absolute size of adjustments falls, and the distribution of price changes becomes more concentrated. Its interquartile range reaches a minimum near the changeover and subsequently recovers. These movements are visible in individual-country time series and are broad-based across product sectors. Event-study estimates that combine the country-sector data show the same pattern, with a pronounced increase in adjustment frequency and a decline in the absolute size of price changes in the changeover month after controlling for country-sector heterogeneity and seasonality.

The Croatian and Bulgarian data provide a more precise view of adjustment timing and show that price changes are concentrated around the conversion date, consistent with the evidence from the historical episodes. However, both changeovers were subject to regulatory monitoring and restrictions on *price increases*. These features make the recent episodes less suitable for identifying firms' unconstrained incentives to adjust prices and highlight the value of the historical CPI data, which form the basis of our main analysis.

We next assess the model more directly using cross-sectional and time-series evidence. The model predicts that price changes at the changeover should be smaller than in steady state and that their relationship with the sector's steady-state adjustment size should be increasing and concave. This pattern appears across country-sector cells. The model also links the compression in adjustment size before and during the changeover to the sector's steady-state frequency of adjustment, and this relationship is also present in the data. Finally, we calibrate the model to the steady-state frequency and size of price changes and choose the single additional changeover parameter to match the increase in adjustment frequency at conversion. The model also reproduces the decline in adjustment size and its subsequent recovery, neither of which is targeted in the calibration.

Related Literature

Our work builds on a large literature on inflation and price setting during euro adoption. Early studies focused on specific product categories, especially restaurant prices, to explain temporary increases in inflation around the conversion (Gaiotti and Lippi, 2004; Hobijn, Ravenna and Tambalotti, 2006; Berardi, Eife and Gautier, 2013). Later work used CPI micro data from individual countries, including Estonia, Austria, and Finland (Meriküll and Rõõm, 2015; ?; Kurri, 2007), while cross-country studies relied on harmonized price indices (Hüfner and Koske, 2008; Ercolani and Dutta, 2006). These papers generally found limited effects on aggregate inflation. Our focus is different. We study the joint behavior of the frequency, size, and dispersion of price changes and use these moments to distinguish among mechanisms of price rigidity. Combining historical CPI micro data from several countries with high-frequency data from recent changeovers allows us to examine whether the same patterns appear across episodes and data sources.

Our paper is closest to Hobijn, Ravenna and Tambalotti (2006) and builds on their insight that the timing of euro adoption is state dependent. They use this mechanism to explain the increase in restaurant-price inflation after the 2002 changeover: price increases that would otherwise have been staggered become concentrated around the conversion, and a "horizon effect" tied to expected marginal costs adds to the increase. This horizon effect requires positive trend inflation to operate and is absent in our framework by design: our model assumes zero drift, so it isolates the pure redenomination channel from any inflation-driven front-loading. The two papers are therefore complementary rather than substitutes. Beyond this theoretical distinction, the empirical targets differ substantially: Hobijn et al. match the aggregate inflation spike in the restaurant sector, while we exploit economy-wide micro data from several episodes to study the joint response of the frequency, size, and dispersion of individual price changes, including both increases and decreases. Our model follows the full transition of the adjustment boundaries and price-gap distributions. This allows us to use the changeover to distinguish among price-setting mechanisms, rather than primarily to account for a temporary increase in inflation. In this respect, the two papers use the same episode to answer different questions: theirs is about the level of inflation, ours is about the distribution of price changes.

The equivalence results discussed above—Alvarez, Lippi and Passadore (2016b) for small monetary shocks, Auclert, Rigato, Rognlie and Straub (2024) for first-order inflation dynamics—are what make an episode of this kind necessary: they imply that aggregate data cannot separate the two classes of model, while the changeover can. The framework may also be useful beyond this application. Similar nonstationarities arise wherever a policy change is announced in advance, such as a pre-announced VAT increase (Karadi and Reiff, 2019). More generally, our results speak to whether price flexibility itself is state dependent, a question that has become especially relevant after the recent return of high inflation. Cavallo, Lippi and Miyahara (2024) document a sizeable increase in the frequency of price adjustments following the large energy shocks of 2022 and use a tractable menu-cost model to show that this frequency response reflects the state dependence of firms' decisions: when markups are sufficiently misaligned, prices are adjusted rapidly, generating a swift pass-through from costs to prices that time-dependent models such as Calvo miss entirely. Alvarez, Gonzalez-Rozada, Neumeyer and Beraja (2011) show that the frequency of price adjustment is largely insensitive to inflation when inflation is low, but rises substantially at higher inflation rates. More recently, Blanco, Boar, Jones and Midrigan (2024) show that this extensive-margin response is quantitatively important for inflation dynamics. The changeover provides a different and complementary way to isolate this margin. It changes the payoff from adjusting prices while leaving desired relative prices unchanged, so any increase in adjustment frequency must reflect the contraction of the inaction region rather than a change in desired markups. As the inaction region contracts, firms with smaller price gaps begin to adjust, so adjustment becomes more frequent even as the average size of price changes falls. The evidence therefore provides a direct test of the state dependence in price adjustment that becomes particularly important in high-inflation environments.

The remainder of the paper proceeds as follows. Section 2 compares the predictions of menu-cost, Calvo, and observation-cost models. Section 3 develops the model and derives its implications for adjustment boundaries, price-gap distributions, and pricing moments. Section 4 describes the data and presents the empirical evidence. Section 5 compares the model's predictions with the cross-sectional and time-series patterns in the data. Section 6 concludes.

2 Menu Costs and Time-Dependent Models

The euro changeover is a purely nominal event: it changes the unit of account but leaves all real fundamentals—preferences, technologies, and the distribution of desired markups—completely unaffected. This makes it an ideal experiment for discriminating among competing models of price stickiness, each of which generates infrequent or dispersed price changes through a different mechanism, and each of which makes a distinct prediction for how firms should behave around a known, purely nominal reform. We describe two such classes of models and their predictions before developing our own framework in Section 3.

2.1 Menu-Cost Model

We begin by describing the steady-state environment in which firms set prices prior to the announcement of the changeover. Firms face a standard menu-cost problem,[2] in which changing the price requires paying a fixed cost ψ, while maintaining a misaligned price entails a

[2]See, e.g., Barro (1972), Dixit (1991), and Golosov and Lucas (2007).

quadratic flow loss Cx^2, where x denotes the gap between the posted price and the firm's ideal (or desired) price. The desired price follows a Brownian motion, so the price gap evolves according to $dx = \sigma\, dW$. Let $\tilde{u}(x)$ denote the firm's value function in steady state. When the firm does not adjust, the value function solves the following Hamilton-Jacobi-Bellman (HJB) equation

$$\rho\tilde{u}(x) \;=\; Cx^2 + \frac{\sigma^2}{2}\tilde{u}_{xx}(x) \qquad \text{for } -\bar{x} < x < \bar{x}.$$

If the firm adjusts, it pays the menu cost ψ and resets the price gap to zero, yielding the (minimized) value $\psi + \tilde{u}(0)$. The boundary conditions are $\tilde{u}(\pm\bar{x}) = \psi + \tilde{u}(0)$, together with the smooth-pasting conditions $\tilde{u}_x(\pm\bar{x}) = 0$.

The solution $\tilde{u}(x)$ is symmetric around zero and strictly convex, with a minimum at $x = 0$. The optimal threshold $\bar{x}$ determines the firm's policy: if $|x| < \bar{x}$, the firm keeps its price unchanged; if $|x| \geq \bar{x}$, it pays the menu cost and resets the gap to zero. In steady state, the threshold $\bar{x}$ is *time-invariant* and fully summarizes the firm's pricing policy.

The threshold $\bar{x}$ also determines two key empirical moments of the price-setting distribution. The size of price adjustments is given by the distance to the boundary, and the frequency of price adjustment is inversely related to the square of that distance. These moments can be expressed explicitly as

$$\bar{x} = \left(\frac{6\psi\sigma^2}{C}\right)^{1/4}, \qquad N = \frac{\sigma^2}{\bar{x}^2}.$$

This canonical menu-cost structure forms the benchmark against which we introduce the time-varying changeover problem in Section 3. The key difference is that, while $\bar{x}$ is constant in steady state, forward-looking firms know the exact date on which prices must be set in euros. The existence of this predetermined implementation date makes the optimal boundaries time-dependent: as the changeover approaches, firms anticipate the forced relabeling and optimally adjust their policies. This forward-looking behavior induces predictable dynamics in adjustment frequencies and price-change sizes.

Because adjustment is costly, firms may also use the changeover date as an opportunity to reconsider their pricing decisions. The changeover forces firms to relabel all prices regardless of the current gap x, and once a firm must "touch" the price tag to convert it into euros, the incremental cost of updating the real price is small relative to paying the menu cost at an arbitrary time. The changeover therefore acts as a coordinated moment of real price reoptimization, generating systematic movements in pricing moments as the mandated date approaches.

2.2 Time-Dependent Models

A second class of models generates infrequent price changes through a timing rule that depends on the passage of time rather than on the state of the price gap. Despite differing in their microfoundations, all models in this class share a common prediction for the changeover: the event is neutral for pricing moments.

Calvo model. The simplest time-dependent model is that of Calvo (1983), in which each firm receives an adjustment opportunity with probability α per period, independently of its price gap. Since the timing of adjustment is exogenous and memoryless, the changeover is a non-event in this framework. Firms that happen to receive an adjustment opportunity at T redenominate their price, but they do so at the same rate and with the same size as in any other period. There is no bunching of adjustments around T, no compression of price-change sizes, and no change in the distribution of price gaps. The changeover is observationally equivalent to any other period.

Observation-cost model. A richer microfoundation for time-dependent pricing is provided by models in which firms face information-processing frictions (Caballero, 1989; Bonomo and Carvalho, 2004; Reis, 2006). In this setting, firms may adjust their posted prices at any instant, but observing the desired (or optimal) price $p^*(t)$ requires paying a fixed cost θ. Let τ denote the time elapsed since the last observation. Between observations, the desired price follows a Brownian motion, $dp^*(t) = \sigma\, dW$, so uncertainty about the firm's ideal price grows linearly with τ.

Let $\tilde{u}(\tau)$ denote the firm's continuation value as a function of the time since its last observation. Conditional on not observing, the firm chooses a price p to minimize the expected quadratic loss $C\,\mathbb{E}\big[(p - p^*(\tau))^2 \,\big|\, p^*(0)\big]$. Since $p^*(t)$ is a martingale, the optimal forecast is $p = \mathbb{E}[p^*(\tau) \mid p^*(0)]$, which implies $\mathbb{E}\big[(p - p^*(\tau))^2 \,\big|\, p^*(0)\big] = \sigma^2\tau$. In the inaction region, the value function solves

$$\rho\tilde{u}(\tau) = C\sigma^2\tau + \tilde{u}'(\tau), \qquad 0 < \tau < \bar{\tau},$$

where $\bar{\tau}$ is the optimal observation threshold. If the firm chooses to observe, it pays θ and resets τ to zero, yielding the boundary condition $\tilde{u}(\bar{\tau}) = \theta + \tilde{u}(0)$. The smooth-pasting condition $\tilde{u}'(\bar{\tau}) = 0$ characterizes the optimal observation time.

The resulting policy rule is simple: the firm observes (and then adjusts its price optimally) whenever $\tau \geq \bar{\tau}$. Because firms can freely reset their posted price whenever they observe, the size of price changes is determined by the forecast error that accumulates between observations, while the frequency of price changes corresponds to the optimal observation rate

$1/\bar{\tau}$.

In this framework, the changeover date does *not* affect the inaction region. Acquiring information about the desired price is costly, but relabeling a posted price is not, so firms have no reason to observe at the moment of the changeover. They simply convert their existing prices into euros without updating their beliefs about the desired markup. As a result, the changeover induces only nominal relabeling, with no change in the underlying pricing moments.

Common prediction. Despite their different microfoundations, Calvo and observation-cost models share a single prediction for the changeover: pricing moments are unaffected. In Calvo, adjustment timing is exogenous and independent of the price gap; in the observation-cost model, relabeling is free so information acquisition is unaffected by the mandated redenomination. In both cases, the changeover is a non-event for the frequency, size, and dispersion of price changes.

2.3 Discussion

Menu-cost and time-dependent models calibrated to match the same steady-state pricing moments generate similar real effects in response to aggregate nominal shocks (Alvarez et al., 2016a,b). The euro changeover breaks this equivalence. In the menu-cost model, the required relabeling lowers the effective marginal cost of adjusting, making the changeover a moment of genuine reoptimization; forward-looking firms anticipate the implementation date and adjust their policies accordingly, generating predictable movements in the frequency and size of price changes. In time-dependent models, by contrast, the changeover should have no real effects, since neither the exogenous adjustment clock in Calvo nor the information-gathering decision in observation-cost models is affected by a costless relabeling requirement.

These contrasting predictions map directly into observable moments. A sharp increase in the *frequency* of price adjustments around the changeover is inconsistent with time-dependent models, which predict neutrality. A simultaneous compression in the *size* and *dispersion* of price changes further sharpens the test: as the inaction region contracts in the menu-cost model, firms with smaller price gaps begin to adjust, so adjustments become more frequent while individual price changes become smaller and more concentrated. As we document in Section 4, all three patterns—higher frequency, smaller size, and compressed dispersion—occur together around the changeover in every country we study. This combination is consistent only with state-dependent menu-cost frictions.

At the same time, the fact that firms are forward-looking and fully aware of the changeover date also explains why a simpler model with an exogenously *mandated* price change at T

would produce the opposite prediction. If all firms were mechanically required to change prices at a fixed future moment, a firm near its usual adjustment boundary would have an incentive to wait and combine its next regular price change with the required changeover. Anticipation alone would therefore *widen* the inaction region and *reduce* adjustment before T—the opposite of what we observe. The data thus point specifically to a model in which the changeover lowers the cost of voluntary adjustment, rather than one in which it mandates it.

3 A Model of Price Setting with a Changeover

In this section we develop a tractable Ss price setting model with a known changeover date. We start by describing the price setting model in a steady state, before the changeover is announced. We study how the announcement, at date $t = 0$, of the changeover at a future date $T > 0$, affects price setting behavior taking into account the costs of pricing in different currencies, as described below.

Steady state. We consider a firm that minimizes the expected discounted value of the costs originated by suboptimal prices. Let x be the (log) difference between the current price and the static profit maximizing price, which we refer to as the price gap. Let Cx^2 be the forgone profits due to the suboptimal price $x \neq 0$, where the constant C converts "squared price gaps" into units of forgone profits per unit of time. With a CES demand system with price elasticity η then $C = \frac{\eta(\eta-1)}{2}$.[3] We assume that in between price adjustments the price gap evolves according to $dx = \sigma\, dW$, where W is a standard Brownian motion. The lack of drift corresponds to an economy with zero inflation, and a firm whose costs are affected by quality shocks. At any moment of time the firm can change its price by paying a fixed cost ψ. We assume that costs are discounted at rate ρ. Given the symmetry of the problem, and the lack of drift, the decision rules are described by a symmetric Ss band, corresponding to an inaction interval $(-\bar{x}, \bar{x})$ and the optimal return point $x^* = 0$. This means that price changes are either $\pm\bar{x}$. A simple computation gives that the expected number of price changes per unit of time is $N = \sigma^2/\bar{x}^2$. We will treat $\{\bar{x}, N\}$ as the two observable statistics from the model. For a given discount rate ρ one can find the ψ/C and σ^2 to rationalize the two observables $\{\bar{x}, N\}$.

The Changeover. The modeling of the changeover is intentionally stylized to preserve analytical tractability and to highlight the core economic mechanism at play.[4] Let $t =$

[3]See proposition 1 in Alvarez, Lippi and Souganidis (2023) where the constant C is denoted B.

[4]Appendix D reports the country-specific timing of dual pricing, cash conversion, and dual circulation. These arrangements differed across episodes and are not modeled separately.

0 denote the announcement date and let T denote the cash-conversion date, when euro banknotes and coins enter circulation and the euro becomes the focal denomination for posted prices. Following the announcement, whenever a firm adjusts its price, it also chooses whether to set the underlying price in euros or in the legacy currency, which we call the "kuna" throughout as a shorthand for the pre-changeover currency of any adopting country; nothing in the analysis is specific to Croatia. We assume that changing this denomination requires paying the menu cost ψ.[5] Redenomination therefore occurs together with a price change. Using the euro before T, or retaining the legacy denomination after T, entails a flow cost ϕ. Because the changeover date is known in advance, the firm's value function becomes explicitly time-dependent, with the optimal policy governed by the remaining distance to T. This section formulates the firm's dynamic problem and characterizes the time-varying decision rules that arise in the presence of a mandated currency switch. We note that both ψ and ϕ are also expressed in units of forgone profits per unit of time.[6]

3.1 Euro Pricing Before and After the Changeover

Let $\tilde{u}^i(\cdot)$ denote the value function for a firm pricing in euros, where $i \in \{A, B\}$ indexes whether the time is After or Before the changeover date T.

Euro Value Function After T. After the cash-conversion date T, setting the underlying price in euros involves no currency-mismatch cost. For a price gap $x \in (-\bar{x}, \bar{x})$, defined as the log difference between the actual and the static profit maximizing price, the value function satisfies the HJB equation

$$\rho \tilde{u}^A(x) = Cx^2 + \frac{\sigma^2}{2}\tilde{u}^A_{xx}(x),$$

where the flow cost Cx^2 is a second order approximation of forgone profits and C is the constant related to the price elasticity of demand discussed above. The value function has boundaries $\tilde{u}^A(\bar{x}) = \psi + \tilde{u}^A(0)$ (value matching) and $\tilde{u}^A_x(\bar{x}) = 0$ (smooth pasting). The solution is

$$\tilde{u}^A(x) = C\left(\frac{x^2}{\rho} + \frac{\sigma^2}{\rho^2} - \frac{2\bar{x}\cosh(\xi x)}{\rho\xi\sinh(\xi\bar{x})}\right) \qquad \text{where } \xi \equiv \sqrt{\frac{2\rho}{\sigma^2}}. \tag{1}$$

[5] We abstract from the simultaneous display of both currencies during dual-pricing periods. Because the conversion rate was fixed, the two displayed amounts were mechanically linked rather than independently chosen. The currency state in the model captures the denomination in which the firm makes its pricing decision.

[6] Both ψ and ϕ are expressed as fractions of annual profits. The menu cost ψ is a one-time payment, so a value of $\psi = 0.01$ means the firm pays 1% of its annual profits each time it adjusts. The flow cost ϕ accrues continuously, so a value of $\phi = 0.01$ means the firm pays 1% of annual profits per year while pricing in the wrong currency. Hence both parameters are directly comparable in magnitude.

This steady-state value function is plotted in Figure 1 (dashed line), as the benchmark prior to the changeover announcement (at time $t = 0$). It also characterizes kuna pricing for $t < 0$, before the changeover is announced: in both cases the firm faces the same stationary problem with no anticipated future cost, and the optimal threshold is $\bar{x}$ in both cases.

We note that we have written the value function $\tilde{u}^A(x)$ in terms of "observables", i.e. in terms of $\bar{x}$ which is the absolute value of the size of price changes, as well as parameters σ^2, C and ρ. For any $\bar{x}$ and σ^2, C, ρ we can find a ψ that rationalizes that $\bar{x}$. To do so, use the expression for $\tilde{u}^A$ in equation (1) and let $\psi = \tilde{u}^A(\bar{x}) - \tilde{u}^A(0)$. Recall that, as can be seen from these equations, the value of $\bar{x}$ depends only on the ratio ψ/C. The average steady state number of price changes is $N = \sigma^2/\bar{x}^2$ that can be used to express σ^2 in terms of N and $\bar{x}$.

Euro Value Function Before T. Before the changeover, i.e. for $t < T$, pricing in euros incurs the per-period cost ϕ.[7] The value function satisfies

$$\rho \tilde{u}^B(x,t) = \phi + Cx^2 + \frac{\sigma^2}{2}\tilde{u}^B_{xx}(x,t) + \tilde{u}^B_t(x,t),$$

with terminal condition $\tilde{u}^B(x,T) = \tilde{u}^A(x)$. Notice how the HJB now takes the form of a partial differential equation (PDE) because the expected value of the costs also depends on the time until the changeover. The solution is

$$\tilde{u}^B(x,t) = \tilde{u}^A(x) + \frac{\phi\left(1 - e^{-\rho(T-t)}\right)}{\rho}. \tag{2}$$

In words, using the euro before T adds a cost on top of the stationary value function $\tilde{u}^A(x)$ that vanishes as $t \to T$.

3.2 Kuna Pricing Before and After the Changeover

This section studies the firm's value functions for kuna-pricing before and after T, and derives the perturbation equations whose solution is characterized in Section 3.3.

Kuna Value Function After T. After the changeover, pricing in kuna incurs the per-unit-time cost ϕ. Compared to the steady state problem, this changes the flow cost of the firm but the problem remains stationary (independent of t), and the value function solves

[7] For example, the *Wall Street Journal* reported concerns among retailers that switching to euro prices too early could deter customers. As one proprietor in Brussels explained: "Converting your prices too early tends to alienate customers. They feel more comfortable with the Belgian franc and may go next door where they know how much it costs." See https://www.wsj.com/articles/SB1005687086876040520.

the ODE (indexed by the flow cost ϕ)

$$\rho u^A(x,\phi) = \phi + Cx^2 + \frac{\sigma^2}{2} u^A_{xx}(x,\phi), \qquad x \in (-\bar{x}^A, \bar{x}^A),$$

with boundaries

$$u^A(\bar{x}^A,\phi) = \psi + \tilde{u}^A(0), \qquad u^A_x(\bar{x}^A,\phi) = 0.$$

The threshold satisfies $\bar{x}^A < \bar{x}$, since pricing in kuna is strictly worse than pricing in euros after T (i.e. since for $\phi > 0$ we have that $u^A(x,\phi) > \tilde{u}^A(x)$). This boundary implies that upon the first price change after T, it is profitable for the firm to adopt the euro and stop paying the penalty ϕ. The optimal decision rule, characterized by the threshold $\bar{x}^A$, is therefore smaller than the steady-state threshold $\bar{x}$, reflecting the additional advantage of adjusting when the firm continues to price in kuna for $t > T$. Using value matching at $\bar{x}^A$ we have

$$u^A(x,\phi) = \frac{\tilde{u}^A(\bar{x}) - \frac{\sigma^2 C}{\rho^2} - \frac{C(\bar{x}^A)^2+\phi}{\rho}}{\cosh(\xi \bar{x}^A)} \cosh(\xi x) + \frac{\sigma^2 C}{\rho^2} + \frac{Cx^2+\phi}{\rho} \quad \text{for } x \in (-\bar{x}^A, \bar{x}^A), \tag{3}$$

where we use that $\tilde{u}^A(\bar{x}) = \psi + \tilde{u}^A(0)$, and that $\bar{x}^A$ solves $u^A_x(\bar{x}^A,\phi) = 0$ or

$$u^A_x(\bar{x}^A,\phi) = \frac{\tilde{u}^A(\bar{x}) - \frac{\sigma^2 C}{\rho^2} - \frac{C(\bar{x}^A)^2+\phi}{\rho}}{\cosh(\xi \bar{x}^A)} \xi \sinh\left(\xi \bar{x}^A\right) + \frac{2C\bar{x}^A}{\rho} = 0 \ .$$

Thus we have

$$u^A(x,\phi) = \begin{cases} \text{given by (3)} & \text{for } x \in (-\bar{x}^A, \bar{x}^A) \\ \tilde{u}^A(\bar{x}) & \text{otherwise} \end{cases}$$

These expressions can be used to verify that if C, ψ, ϕ are multiplied by a common positive constant, then both $\bar{x}$ and $\bar{x}^A$ remain the same. In words, the optimal decision depends on the cost relative to benefits, i.e. on the ratios $\psi/C, \phi/C$. As in the case of $\tilde{u}^A$, both the value function and the threshold are written in terms of observables –see the discussion above.

Kuna Value Function Before T. Following the announcement at $t = 0$, the present value cost of being a kuna pricer increases relative to the stationary benchmark $\tilde{u}(x)$: the firm anticipates that at some point it will incur costs from currency mismatch, raising its continuation value of remaining in kuna. For all $t \in (0,T)$, the kuna value function $u^B(x,t,\phi)$ satisfies the PDE

$$\rho u^B(x,t,\phi) = Cx^2 + u^B_{xx}(x,t,\phi)\frac{\sigma^2}{2} + u^B_t(x,t,\phi) \qquad \text{for} \quad x \in \Big(-\bar{x}(t), \bar{x}(t)\Big), \tag{4}$$

with terminal condition $u^B(x,T,\phi) = u^A(x,\phi)$.

Let $T_0 > 0$ denote the time when euro adoption becomes optimal upon adjustment, characterized below. For $t \in (0, T_0)$, a firm that adjusts resets to $x = 0$ and stays in kuna; the value matching and smooth-pasting conditions at the boundary therefore read

$$u^B(\bar{x}(t), t, \phi) = \psi + u^B(0, t, \phi) \qquad , \qquad u^B_x(\bar{x}(t), t, \phi) = 0. \tag{5}$$

The changeover announcement raises the firm's continuation value, because of the presence of the currency mismatch costs.

After T_0, euro adoption becomes optimal upon adjustment and the threshold $\bar{x}(t)$ begins to contract. The boundary conditions in equation (4) now read

$$u^B(\bar{x}(t), t, \phi) = \psi + \tilde{u}^B(0, t, \phi) \qquad , \qquad u^B_x(\bar{x}(t), t, \phi) = 0, \tag{6}$$

with the terminal boundary condition at $t = T$

$$u^B(x, T, \phi) = u^A(x, \phi) \qquad \forall x. \tag{7}$$

Perturbation for Kuna Pricing Before T. The PDE in equation (4) with time-varying boundaries $\underline{x}(t)$ and $\bar{x}(t)$ does not admit a closed-form solution in general. We therefore proceed by linearizing the problem around $\phi = 0$, to obtain an analytical characterization of how the boundaries and value function respond to the introduction of a (small) penalty.

Define the perturbation for $u^B(x, t, \phi)$ and $\bar{x}(t, \phi)$ as

$$v^B(x,t) \equiv \frac{\partial u^B(x,t,\phi)}{\partial \phi}\Big|_{\phi=0} \quad , \quad \alpha(t) \equiv \frac{\partial \bar{x}(t,\phi)}{\partial \phi}\Big|_{\phi=0}. \tag{8}$$

Differentiating equation (4) with respect to ϕ gives the PDE that v^B solves

$$\rho v^B(x,t) = v^B_{xx}(x,t)\frac{\sigma^2}{2} + v^B_t(x,t) \quad \text{for} \quad x \in (-\bar{x}, \bar{x}) \,. \tag{9}$$

The boundaries at $x = \pm\bar{x}$ are obtained by differentiating equation (6) and equation (2)

$$v^B(-\bar{x}, t) = v^B(\bar{x}, t) = \frac{1 - e^{-\rho(T-t)}}{\rho} \tag{10}$$

with terminal condition obtained by differentiating equation (7):

$$v^B(x, T) = v^A(x) \quad \forall x, \tag{11}$$

where $v^A(x) \equiv \frac{\partial u^A(x,\phi)}{\partial \phi}\big|_{\phi=0}$ is the perturbation of the kuna-after-T value function. Since we have the closed-form solution for $u^A(x,\phi)$, we can compute v^A directly:

$$v^A(x) = \frac{1}{\rho}\left(1 - \frac{\cosh(\xi x)}{\cosh(\xi \bar{x})}\right), \qquad \xi \equiv \sqrt{\frac{2\rho}{\sigma^2}}.$$

This expression serves as the terminal condition in equation (11).[8] It is obtained by differentiating equation (3) with respect to ϕ keeping $\bar{x}^A$ fixed (this partial derivative gives the correct expression for $v^A(x)$ by an application of the envelope theorem at each x).

Likewise the perturbation of the smooth pasting condition in equation (6) gives

$$u^B_{xx}(\bar{x},t,0)\ \alpha(t) + v^B_x(\bar{x},t) = 0$$

where $u^B_{xx}(\bar{x},t,0) = \tilde{u}^A_{xx}(x)$ since it is evaluated at $\phi = 0$. This condition will be used to study the evolution of the decision rule $\bar{x}(t) = \bar{x} + \phi\,\alpha(t)$.

The First Adoption Time: T_0. The model identifies the first time at which it is optimal for a firm to adopt the euro before T. Adoption occurs at the earliest T_0 such that

$$u^B(0,T_0) = \tilde{u}^B(0,T_0),$$

where $\tilde{u}^B$ was given in equation (2). For $t < T_0$, firms optimally remain in kuna when adjusting; for $t > T_0$, every adjustment results in euro adoption. The value of adjusting at each t, conditional on staying in kuna or adopting the euro early, is illustrated in Figure A1: the two value curves intersect at T_0, marking the onset of the early euro adoption.

3.3 Analytical Solution: Dynamics of the Decision Rule

To characterize the time path of $\bar{x}(t)$ analytically, we solve the perturbation problem defined in the previous section in closed form. The key step is to map the heat equation (9), with the Dirichlet boundary conditions given in (10), into the general solution provided in Lemma 1 of Appendix B. We have the following

Proposition 1. *Consider the kuna problem before the changeover. Let $N = \sigma^2/\bar{x}^2$ be the steady-state adjustment frequency. We have:*
(i). For $x \in [-\bar{x}, \bar{x}]$ and $t \in [T_0, T]$, the approximate value function is $u^B(x,t) = \tilde{u}^A(x) +$

[8] In particular it is used in the projection coefficients $a_j(T) \equiv \langle \varphi_j, v^A \rangle$ computed in Appendix B.1.

$\phi\, v^B(x,t)$ where

$$v^B(x,t) = \frac{1 - e^{-\rho(T-t)}}{\rho} + \sum_{j=1,3,5,\ldots} b_j(t)\, \sin\left(j\pi\, \frac{\bar{x}+x}{2\bar{x}}\right),$$

where $\xi \equiv \sqrt{2\rho/\sigma^2}$, $\lambda_j \equiv \rho + \frac{N\,(j\pi)^2}{8}$ and

$$b_j(t) \equiv \frac{4}{j\pi}\left[\frac{1}{\rho}\left(1 - \frac{(j\pi)^2}{4(\xi\bar{x})^2 + (j\pi)^2}\right) e^{-\lambda_j(T-t)} - \frac{1 - e^{-\lambda_j(T-t)}}{\lambda_j}\right].$$

(ii) The first adoption time T_0 satisfies $\sum_{j=1,3,5,\ldots} b_j(T_0)\, \sin\left(\frac{j\pi}{2}\right) = 0$.
(iii) The approximate inaction boundary is $\bar{x}(t) = \bar{x} + \phi\, \alpha(t)$, where

$$\alpha(t) = \frac{\sum_{j=1,3,5,\ldots} b_j(t)\, j\pi}{\tilde{u}^A_{xx}(\bar{x})\, 2\bar{x}} \qquad \text{and} \qquad \tilde{u}^A_{xx}(\bar{x}) = \frac{2C}{\rho}\left(1 - \frac{\xi\bar{x}}{\tanh(\xi\bar{x})}\right) < 0\,. \tag{12}$$

This proposition gives a closed form solution for the first order correction to the value function, $v^B(x,t)$, and to the path of the threshold, $\alpha(t)$. As defined above, both approximations are obtained, up to the first order, by perturbing the problem relative to the cost ϕ, see equation (8). As in the previous expressions for $\tilde{u}^A(x)$ and $u^A(x)$, we express the closed form solution in terms of steady state "observables" such as the frequency and size of price changes N and $\bar{x}$ (see the discussion after equation (1)).

The next proposition considers the case of a small discount rate ρ and obtains a sharp characterization of the pricing behavior before the changeover:

Proposition 2. Consider the first adoption time T_0 characterized in result (ii) from Proposition 1. As $\rho \to 0$ then $N(T - T_0)$ is a pure constant (independent of any model parameter), approximated by:

$$T - T_0 \approx T - \tilde{T}_0 = \frac{1}{N}\frac{8\log 3}{\pi^2}. \tag{13}$$

Proposition 2 gives a characterization of the single determinant of T_0 and provides an approximation of its value. Two remarkable properties are revealed by equation (13). The first adoption time relative to T depends only on N. In particular it does not depend on the cost ϕ. The reason for this is that ϕ affects both the cost and benefit of pricing in Euros before the changeover, and to a first order this effect is proportional to both. Second, changing the steady state frequency keeps $(T - T_0)N$ constant. This can be understood by realizing that N measures the relevant units of time. The discount rate ρ is also measured in time units, that is why this sharp result requires $\rho \to 0$. This result has important implications. Sectors with more price changes (higher N) will start adopting the new currency closer to the T threshold. Intuitively, sectors where prices change more frequently face more opportunities to adopt the new currency. A second implication of the approximate expression for T_0 is that the change in behavior induced by the changeover, i.e., the starting time for price changes with re-denomination, is relatively close to T for sectors featuring a high N. For instance, if a sector changes its price twice a year, so $N = 2$, then $T - T_0 \approx 0.44$ of a year or 5.4 months. For a sector with $N = 4$ the time window where behavior is affected shrinks to about 0.22 of a year, or 2.7 months. Thus, in spite of the forward looking nature of the firm problem, changes in frequencies and size of price changes will be concentrated in the last few months before the changeover.

We now turn to the analysis of the dynamics of the threshold $\bar{x}$ for the case of a small discount rate.

Proposition 3. Consider the threshold $\bar{x}(t)$ for the kuna problem before the changeover date, characterized in result (iii) from Proposition 1. Defining $\hat{\alpha}(t) \equiv \alpha(t)C\bar{x}$, as $\rho \to 0$ we have

$$\bar{x}(t) = \bar{x} + \hat{\alpha}(t)\tfrac{\phi}{C}\tfrac{1}{\bar{x}} + o\left(\tfrac{\phi}{C}\right) \quad \text{for} \quad t \in (T_0, T) \tag{14}$$

$$\hat{\alpha}(t) \equiv -6 \sum_{j=1,3,5,\ldots} \frac{3e^{-N(T-t)\frac{(j\pi)^2}{8}} - 1}{(j\pi)^2} \quad \text{for } t \in [\tilde{T}_0, T) \tag{15}$$

$$\hat{\alpha}(\tilde{T}_0) > 0\,, \quad \hat{\alpha}(T) = -\tfrac{3}{2}\,, \quad \frac{\partial \hat{\alpha}(t)}{\partial t} < 0, \left.\frac{\partial \hat{\alpha}(t)}{\partial t}\right|_{t=T} = -\infty \ \text{ and } \ \frac{\partial^2 \hat{\alpha}(t)}{\partial t^2} < 0\,. \tag{16}$$

Proposition 3 delivers the time path of the inaction boundary $\bar{x}(t)$ for a small cost ϕ/C. A few comments are in order. First, and most important, $\alpha(t) < 0$ for an interval (t_0, T) where $T_0 < t_0 < T$, where t_0 satisfies $\alpha(t_0) = 0$. Thus, prior to the changeover, the size of price changes is smaller than in the steady-state. Second, the time path of the gap with respect to the steady state, i.e. $\bar{x}(t) - \bar{x}$, is described by the product of $\phi/(C\bar{x})$ and the function of time $\hat{\alpha}(t)$, which depends only on one parameter, namely N. Furthermore, in this function time is measured relative to the expected duration of a price change in steady state. Moreover, the profile of the function $\hat{\alpha}(t)$ with respect to time is decreasing, sharply so when t is near T. In words, Proposition 2 not only shows that all the redenominations are concentrated in the short time interval (T_0, T), but the concavity of α and its infinite derivative shown in

Proposition 3 imply that most of the adjustment occurs close to the changeover date T. Two more properties are worth emphasizing. First, the size of the price gap of the price changes just before the changeover depends on $\bar{x}$, i.e. $\bar{x}(T) - \bar{x} = -\frac{3}{2}\frac{1}{\bar{x}}\frac{\phi}{C}$. In words, the relationship between $\bar{x}(T)$ and $\bar{x}$ for a fixed ϕ/C is increasing in $\bar{x}$, below the 45 degree line, and concave in $\bar{x}$. Second, the ratio $(\bar{x}(t_2) - \bar{x})/(\bar{x}(t_1) - \bar{x})$ for $T_0 < t_1 < t_2 < T$ depends only on Nt_1 and Nt_2, i.e. it is independent of any other parameters.

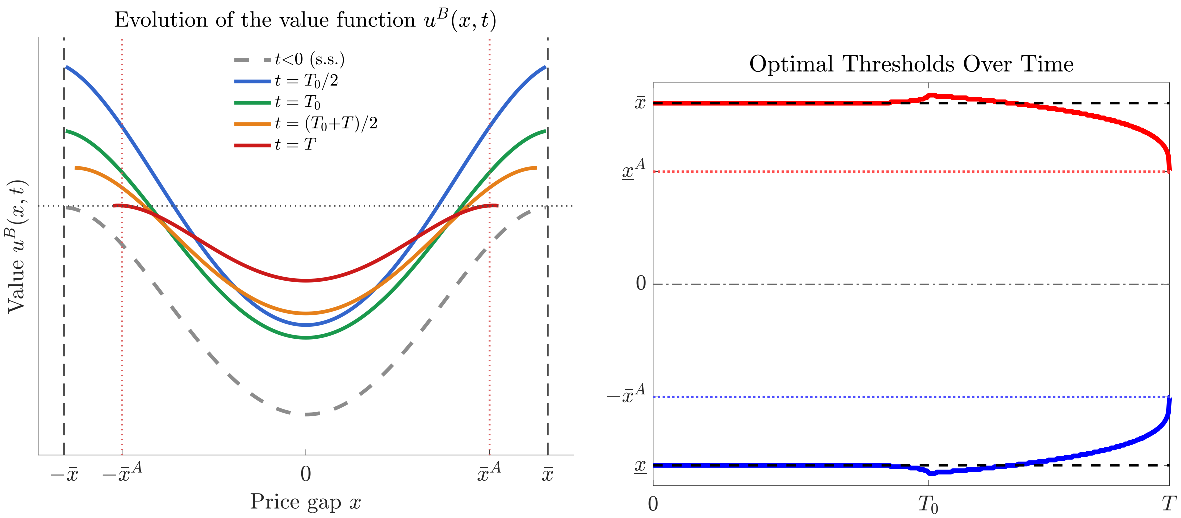
The rise in pricing activity before the changeover is not a mechanical consequence of anticipation. If firms simply knew that they would have to change their price tags at a fixed future date, a firm close to its usual adjustment boundary would have an incentive to wait and combine its next price change with the required changeover. Anticipation alone would therefore widen the inaction region and reduce adjustment before T. This force appears in our model slightly before T_0, when the inaction threshold $\bar{x}(t)$ widens above $\bar{x}$, namely $\alpha(T_0) > 0$. A firm near the boundary has an additional incentive to delay adjustment, since waiting until after T_0 allows it to redenominate into euros upon the next price change.[9] This can be seen in the right panel of Figure 1, which also shows the other properties highlighted by the proposition for $t \in (T_0, T)$.[10]

The perturbation is valid for small ϕ; for larger values the exact boundaries require numerical solution, as do the cross-sectional distributions $k(x,t)$ and $e(x,t)$ discussed in Section 3.4.

[9] For t slightly below T_0 the firm anticipates the benefits of switching to the euro (recall that $\tilde{u}^B(0,t) < u^B(0,t)$ for $t > T_0$). The firm is therefore willing to tolerate a slightly larger price gap and wait a little longer to reap the redenomination benefits.

[10] The figure is obtained by solving the firm problem numerically using finite differences with a very small grid size. The numerical solution shows that the analytical approximations reported in the proposition are accurate.

Figure 1: The kuna Value Function, $u^B(x, t)$ and Threshold Rule $\bar{x}(t)$ Before T



Note. Left Panel: The gray dashed curve shows the pre-announcement steady state $\tilde{u}(x)$. Blue, green, orange, and red curves correspond to $t = T_0/2$, $t = T_0$, $t = (T_0 + T)/2$, and $t = T$ respectively. The horizontal dotted line marks the adjustment trigger level $\psi + \tilde{u}(0)$; vertical dashed lines mark the stationary boundaries $\pm\bar{x}$ and the post-changeover boundaries $\pm\bar{x}^A$. Right panel: The red (blue) curve shows the upper (lower) boundary. Dashed lines mark the steady-state threshold $\pm\bar{x}$; dotted lines mark the post-changeover threshold $\pm\bar{x}^A$ at which kuna-pricing firms eventually adjust after T.

Value Function and Threshold Dynamics. Figure 1 illustrates how the firm's value function $u^B(x, t)$ and the optimal threshold $\bar{x}(t)$ evolve as the changeover approaches. The left panel is constructed using the perturbation solution of Proposition 1, with the level of each curve normalized to satisfy the value matching condition in equation (6) at the time-varying boundary $\bar{x}(t)$. The right panel uses the exact numerical solution.

The left panel overlays the value function $u^B(x, t)$ at five time slices: before the announcement ($t < 0$), u^B coincides with the stationary benchmark $\tilde{u}(x)$: a symmetric U-shape with inaction region $\pm\bar{x}$, shown as the gray dashed curve (this is equivalent to the steady state value function given in equation (1)). Once the changeover is announced and t approaches T_0, the future flow penalty ϕ raises continuation values, but there is little effect on the decision threshold initially since euro adoption is not yet optimal.

The right panel shows the time path of the inaction boundary $\bar{x}(t)$, as characterized by Proposition 1. At $t = T_0$ (green curve), the firm becomes indifferent between remaining in kuna and adopting the euro early. For $T_0 < t < T$, the inaction region given by the interval $(-\bar{x}(t), \bar{x}(t))$, visibly narrows as the remaining time until T shrinks, reducing the cost of early euro adoption and making redenomination increasingly attractive upon adjustment.

As a result, $\bar{x}(t)$ becomes very small at $t = T$, implying that the mass of adjustments occurring exactly at the changeover consists of small price changes.[11]

The shape of $\bar{x}(t)$ depends crucially on whether the flow penalty applies before or only after the changeover. A positive pre-changeover penalty ($\phi^B > 0$) delays the start of euro adoption, so adoption begins only at $T_0 > 0$. In contrast, if $\phi^B = 0$, there is no disadvantage to adopting early, and euro adoption begins immediately at $t = 0$. The post-changeover penalty $\phi^A > 0$ governs the contraction of the boundary at T: it determines how sharply $\bar{x}(t)$ shrinks at the changeover. The case with $\phi^B = 0$ and $\phi^A > 0$ is analytically equivalent to a model with a lump-sum inspection penalty.[12]

3.4 Distributions of Price Gaps

The value functions and optimal thresholds characterized above govern the evolution of the cross-sectional distribution of price gaps. We track two populations: firms still pricing in kuna, with density $k(x,t)$, and firms that have already adopted the euro, with density $e(x,t)$.

For $t < 0$. Before the changeover is announced, the economy is in steady state. All firms price in kuna, the inaction threshold is constant at $\bar{x}$, and the cross-sectional distribution of price gaps is exactly the stationary tent distribution $\tilde{k}(x) = (\bar{x} - |x|)/\bar{x}^2$, which solves the stationary KFE

$$\frac{\sigma^2}{2}\tilde{k}_{xx}(x) = 0, \qquad x \in (-\bar{x}, 0) \cup (0, \bar{x}),$$

with absorbing boundaries at $\pm\bar{x}$, inflow at the reset point $x = 0$, and $\int \tilde{k}(x)\,dx = 1$.

For $t \in (0, T_0)$. After the announcement but before the first adoption time, no firm finds it optimal to adopt the euro early upon adjustment so that the euro distribution is zero, $e(x,t) = 0$. The numerical solution shows that the inaction threshold for the kuna pricing firms remains very close to $\bar{x}$ throughout this period, so the distribution $k(x,t)$ is approximately stationary and well approximated by $\tilde{k}(x)$.

For $t \geq T_0$. At T_0, euro adoption becomes optimal upon adjustment. Firms that reach the boundary $\pm\bar{x}(t)$ now exit permanently into the euro population rather than resetting at $x = 0$ within the kuna distribution. The density $k(x,t)$ therefore satisfies

$$k_t(x,t) = \frac{\sigma^2}{2}k_{xx}(x,t), \qquad x \in (\underline{x}(t),\, \bar{x}(t)),$$

with initial condition $k(x, T_0) \approx \tilde{k}(x)$ and absorbing boundary conditions $k(\underline{x}(t), t) = k(\bar{x}(t), t) =$

[11] Using the stationary distribution of price gaps, an absolute boundary of 15% corresponds to an average adjustment of about 5%, since: Average absolute size $\approx 2\int_0^{\bar{x}} x\,\frac{\bar{x}-x}{\bar{x}^2}\,dx = \frac{\bar{x}}{3}$.

[12] Details of this model are available upon request.

0 and no inflow at the origin. As a result, $\int k(x,t)\,dx$ declines monotonically toward zero as mass is permanently transferred to the euro population.

Firms that have adopted the euro face the stationary problem with fixed thresholds $\pm\bar{x}$. Their density $e(x,t)$ satisfies

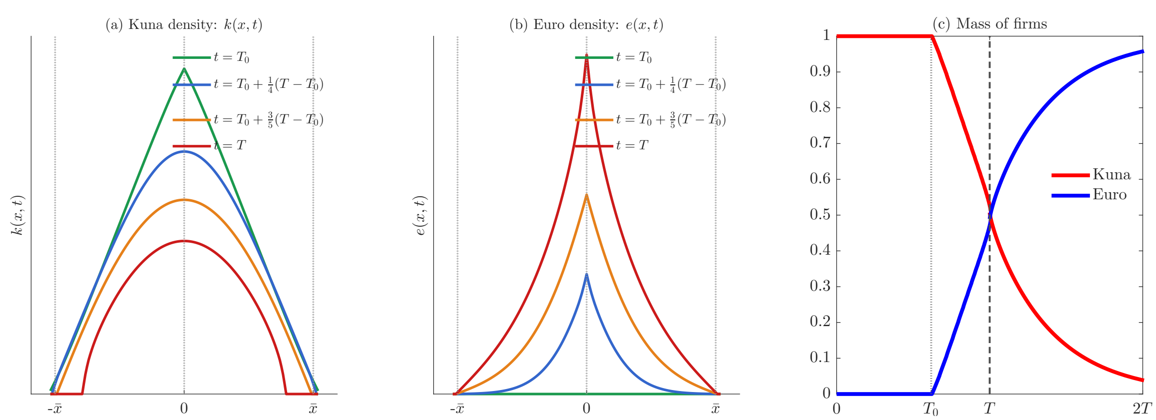
$$e_t(x,t) = \frac{\sigma^2}{2} e_{xx}(x,t), \qquad x \in (-\bar{x}, 0) \cup (0, \bar{x}),$$

with absorbing boundaries $e(-\bar{x},t) = e(\bar{x},t) = 0$, inflow at the origin from euro firms that adjust, and initial condition $e(x,T_0) = 0$. Mass conservation now applies only to the aggregate:

$$\int_{\underline{x}(t)}^{\bar{x}(t)} k(x,t)\,dx \;+\; \int_{-\bar{x}}^{\bar{x}} e(x,t)\,dx \;=\; 1.$$

The evolution of $k(x,t)$ and $e(x,t)$ is illustrated in Figure 2. Panel (a) shows that $k(x,t)$ starts at the steady-state tent distribution at $t = 0$, remains close to it until T_0, then erodes from the boundaries as firms begin adopting the euro early, with a non-negligible mass of firms still pricing in kuna at $t = T$. Panel (b) shows the complementary buildup of $e(x,t)$, which grows from zero at T_0 and converges toward the stationary tent distribution $\tilde{k}(x)$ as $t \to \infty$. Panel (c) summarizes the aggregate dynamics: the kuna mass declines monotonically from T_0 onward while the euro mass grows correspondingly, with the transition accelerating sharply near T.

Figure 2: Distributions of Price Gaps and Mass of Firms



Note: Panel (a) shows the cross-sectional density of price gaps $k(x,t)$ for firms still pricing in kuna at four time slices: $t = T_0$ (green), $t = T_0 + \frac{1}{4}(T - T_0)$ (blue), $t = T_0 + \frac{3}{5}(T - T_0)$ (orange), and $t = T$ (red). Panel (b) shows the corresponding density $e(x,t)$ for firms that have already adopted the euro. Panel (c) plots the total mass of firms pricing in kuna (red) and euros (blue) from T_0 to $2T$; the vertical dashed line marks the changeover date T.

Pricing moments. The two densities deliver the key empirical moments. Let $R(t)$ denote the rate of redenomination—the flow of firms switching from kuna to euro pricing. We could also think of $R(t)$ as the price changes where the initial price is denominated in kuna, for $t \in (T_0, T)$. On the other hand $N(t)$ denotes the total number of price changes. These two rates obey:

$$R(t) = \frac{\sigma^2}{2}\left[k_x(\underline{x}(t), t) - k_x(\bar{x}(t), t)\right] > 0,$$

$$N(t) = R(t) + \frac{\sigma^2}{2}\left[e_x(-\bar{x}, t) - e_x(\bar{x}, t)\right] > 0 \ .$$

We are interested in the ratio $r(t) = R(t)/N(t)$, measuring what fraction of the price changes corresponds to redenomination, or equivalently the fraction of price changes that start with a price in kuna. Note that $t(T_0) = 1$, since all price changes at that point originate from kuna prices. Notice moreover that if $\phi = 0$, then $r(t) = 1$ for all t. By a change of variables one can show that $r(t)$ does not depend on $\bar{x}$ and that it only depends on $t \cdot N$, as was the case for $\hat{\alpha}$.

The average absolute size of price changes for $t \in (T_0, T)$ is $S(t) = \bar{x}(t) \cdot r(t) + \bar{x} \cdot (1 - r(t))$, a weighted average between the steady-state $\bar{x}$, for the firms that have already redenominated, and the smaller price changes $\bar{x}(t)$ for the firms switching from kuna to euro. Using the same expansion used before we write $r(t) = 1 + \hat{r}(t)\phi + o(\phi)$. Thus

$$S(t) - \bar{x} = (\bar{x}(t) - \bar{x}) \cdot r(t) = \alpha(t)\phi + o(\phi) \tag{17}$$

Figure 1 (right panel) shows that $\bar{x}(t)$ falls as the changeover approaches, hence for a small ϕ the model predicts that the gap with the average size is governed by the $\alpha(t)$ characterized in Proposition 3.

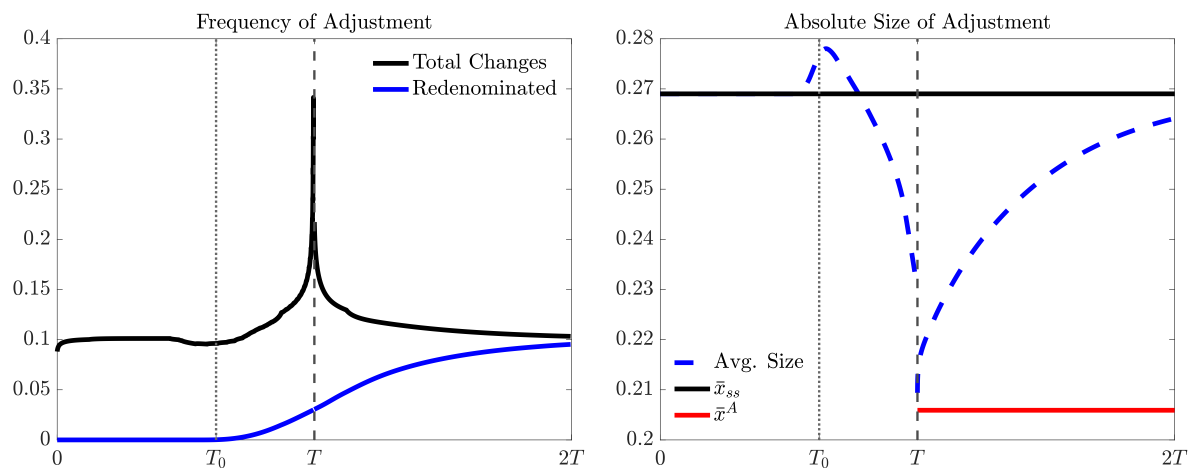
3.5 Model Predictions for Observable Moments

This section derives the model's predictions for the observable moments of the price-change distribution — the frequency and average absolute size of price adjustments — as functions of time relative to the changeover date. These are the objects we take to the data in Section 4, and the model's key predictions are a spike in frequency and a compression in size as $t \to T$.

Figure 3 summarizes the implications of the model for the frequency and the size of price adjustments around the changeover. The left panel plots the total adjustment rate and the subset of adjustments associated with euro adoption. For $0 < t < T_0$, all adjustments correspond to price changes in kuna and the total frequency remains close to its steady-state level. At T_0, early euro adoption begins, and the frequency of redenomination starts to rise. As t approaches T, the shrinking of the kuna inaction region sharply increases the probability that a firm hits the adjustment boundary, generating a pronounced spike in the total adjustment rate. For $t > T$, the frequency gradually returns toward its post-changeover steady state as the economy converges to the euro-only pricing regime.

The right panel shows the evolution of the absolute size of price changes. Before the announcement the absolute size remains at its steady-state benchmark. Once early adoption begins at T_0, the boundary $\bar{x}(t)$ starts to move inward, and the average size of price changes declines. This decline becomes steep as t approaches T, reflecting the rapid contraction of the inaction region predicted by the model, so the mass of adjustments executed around T consists of small price changes on average. For $t > T$, as the economy converges to the euro-only regime, adjustment sizes rise gradually toward their steady state level.

Figure 3: Optimal Pricing Before and After Changeover



Note: The left panel plots the frequency of price adjustments predicted by the model. The right panel shows the corresponding absolute size of adjustments.

The model delivers sharp predictions for how pricing behavior should evolve before, during, and after the changeover. We now turn to the data to evaluate these predictions.

4 Micro Data from Six Euro Changeovers

We use data from six countries that experienced the euro changeover: Austria, Finland, Greece, Slovakia, Croatia, and Bulgaria. Studying these episodes at the micro level is difficult because doing so requires access to the underlying price quotes, and for the earlier changeovers these data are confidential and in many cases are no longer available. For Austria, Finland, Greece, and Slovakia, we were able to obtain historical CPI micro data collected by the respective statistical offices.[13] Access to these data outside the statistical agencies is rare. With the exception of Finland, where the data can only be accessed on site, we were able to work with the data remotely, allowing us to combine the historical files in a common panel and estimate regressions across countries and sectors. Bringing together these independently collected CPI micro datasets also allows us to examine whether the same pricing patterns appear across different statistical systems and changeover episodes. The datasets complement one another because of differences in country size, data coverage, and the approach to the euro transition. Austria, Finland, and Greece followed the Madrid-style approach, with a transitional period in which the euro existed as book money before euro cash was introduced. Slovakia, in contrast, followed a Big-bang approach, in which the euro became the official currency and euro cash was introduced at the same time, with only a short period of dual circulation (see Appendix D for more details).

Additionally, we collected daily posted prices for Croatia and Bulgaria, which adopted the euro on January 1, 2023, and January 1, 2026, respectively. Both countries followed relatively fast transitions, with short dual-circulation periods during which the euro and the legacy currency were accepted simultaneously. For Croatia, we collected daily posted prices through web scraping beginning approximately six months before the changeover and continuing for almost two years after it. For Bulgaria, we obtained daily retailer-level data covering more than 200 retailers before and after the conversion. We discuss these two recent changeovers and the associated data in greater detail in Section 4.2.

4.1 CPI Data: Austria, Finland, Greece, Slovakia, and the UK

Austria: We obtained the micro-data underlying the Austrian CPI from 1996 to 2004, covering the years surrounding the euro changeover, which took place in January 2002. The data is collected by Statistics Austria every month and covers approximately 90% of the total CPI (Baumgartner et al., 2005). Each observation includes the following information: product, product category (five-digit ECOICOP), date, outlet, region, price, unit, and indicators for whether the price is missing or whether a product was "on sale" when its price was sampled in a particular month. We use this sales flag to calculate statistics on the frequency and size of price changes, excluding sales. The dataset also contains CPI expenditure weights for each

[13] We inquired about the availability of CPI micro data across statistical agencies and central banks in all countries that experienced the euro changeover. These data are often not disseminated outside the statistical systems of these countries, are no longer technically accessible because of changes in data systems, or have been destroyed. For example, retention periods are typically tied to the period during which previously published results might need to be reconsidered or recalculated.

product category. Overall, the dataset includes a total of 63,267 products (item $\times$ outlet) across 842 product categories and 21 cities.

Finland: We use the micro-data underlying the Finnish CPI from 1997 to 2008. In Finland, the euro changeover took place in January 2002. The data can only be accessed at Statistics Finland and comprise 55% of the items in the CPI basket (Kurri, 2007). Each observation includes information on the product, price, product category (four-digit COICOP), collection method (e.g., interview), geographic area (e.g., south, east), outlet, type of outlet, and currency (one euro is equivalent to 5.94573 markka). The dataset also includes indicators for whether the price is missing, imputed, or whether the product is on sale in a particular month. The number of collected products is approximately 498 items and 2,600 outlets. The coverage of the data varies significantly around the euro changeover. The data include about 50,000 observations per month. Three months before the euro changeover, the number of observations increased to 90,000, and from January 2002 to May 2002, it remained around 65,000 observations before dropping back to 50,000. These changes in coverage occurred for two reasons. First, Statistics Finland collected more price information to measure the CPI as accurately as possible during the euro changeover. Second, before 2012, Statistics Finland resampled products every five years, leading to a large update of products around the euro changeover. As a result, only 30% of the product-outlet pairs observed before the changeover remain in the data afterward. Because changes in sample composition could mechanically affect the pricing moments, we restrict the analysis to product-outlet pairs observed both before and after the changeover.

Greece: The CPI micro-data for Greece cover approximately 95% of the CPI items. We obtained the data from the Hellenic Statistical Authority (ELSTAT) for the years 1998–2005 (the euro changeover took place in January 2002). The dataset contains information on products, product categories (two-digit COICOP), regions (cities or towns), and outlets. However, it does not include information on product substitutions or products on sale. Before 2002, all prices were quoted in drachma (one euro is equivalent to 340.750 drachma). Starting in 2001, there was a significant increase in price collection in cities and outlets, resulting in approximately 80% more observations. In total, the dataset contains information on 726 distinct products, 26 cities, and 103 outlets. However, it does not provide data on product substitutions or sales.

Slovakia: We obtained CPI pricing data for Slovakia from the Slovak Statistical Office (SSO). The data are monthly and cover the period from 2007 to 2012 (the euro changeover took place in January 2009). Each observation includes information on the product, price, product category (five-digit COICOP), period, outlet, and district. The dataset contains prices for 754 products collected across 39 districts and 14,594 outlets. Before 2009, all prices were quoted in korunas (one euro is equivalent to 30.1260 korunas). The data also include an indicator for whether the price is missing or imputed. However, the dataset does not contain information on whether a product was on sale in a particular month.

United Kingdom: We use the CPI micro-data from the UK's Office for National Statistics (ONS). Since the UK did not experience a euro changeover, we use this dataset primarily as a reference. Although the dataset is available from 1996 onward, we focus on the period surrounding the euro changeover. The dataset is publicly available and includes prices for approximately 1,000 individual goods or services each month, collected locally from more than 10,000 retail stores across the UK. It also includes indicators for whether the price is missing, out of stock, or if a product substitution occurred. Additionally, the dataset provides information on whether a product was on sale in a particular month, product categories (two-digit COICOP), and covers about 57% of the UK CPI basket.

Table 1 summarizes our datasets. The main advantage of the Austrian data is that, unlike the Greek data, it includes indicators for sales and product substitutions. Moreover, the dataset was not subject to a large resampling phase around the euro changeover, as was the case with the Finnish data. The Slovak data complement the other datasets, as the country experienced a different changeover approach during a different time period. Only the Slovak and Finnish datasets contain price quotes in their original currencies before the changeover.

Pricing Moments: We begin computing the frequency of price changes in each data set by constructing an indicator that equals one if the price for a given item×store is different than that of the previous month. We exclude product substitutions and imputations whenever that information is available.[14]

An important issue during the euro changeover was rounding, as in many countries, the minimum unit of account in the national currency was smaller than the lowest euro denomination. To address this, the European Commission issued guidelines stating that after

[14] The Austrian data identify sales and product substitutions. To preserve a comparable outcome across countries, the cross-country baseline includes sales because Greece and Slovakia do not contain comparable sales indicators. For Austria we additionally report estimates that exclude sales; the qualitative frequency and size patterns around the changeover are unchanged. Panel (a) of Figure C1 shows that in the Austrian case, we do not find evidence that sales or product substitutions increase around the changeover. Panel (b) shows similar evidence for product substitutions and imputations in Slovakia.

Table 1: Data Description: CPI Price

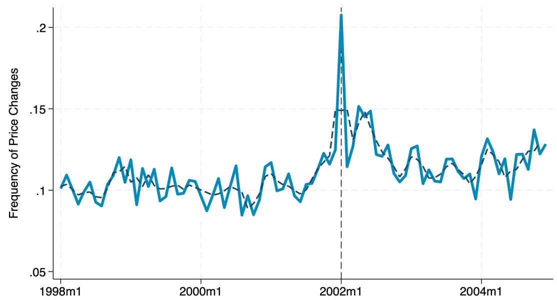
Note: The table summarizes each of the five CPI micro-data sets we use. Since the UK did not experience a changeover, the UK micro-data–which is the only dataset that is publicly available–is primarily used as a reference.

	Austria	Finland	Greece	Slovakia	UK
Dates	1996-2004	1997-2008	1998-2005	2007-2012	1996-2004
Outlets	3,708	2,600	103	14,594	10,665
Item-Outlet	63,267	78,158	42,264	110,369	431,752
COICOP	Class	Subclass	Division	Subclass	Subclass
Sales	✓	✓	✗	✗	✓
Subs.	✓	✓	✗	✓	✓
Imputation	✗	✗	✗	✓	✓
Currency	✗	✓	✗	✓	✓
Obs.	4,086,209	7,274,026	3,171,673	6,213,783	10,573,127
Changeover	Jan 2002	Jan 2002	Jan 2002	Jan 2009	N/A
Approach	Madrid	Madrid	Madrid	Big-bang	N/A
Exchange	13.7603 ATS	5.94573 FIM	340.75 GRD	30.126 SKK	N/A
Access	Restricted Remote	Restricted On-Site	Restricted Remote	Restricted Remote	Public Remote

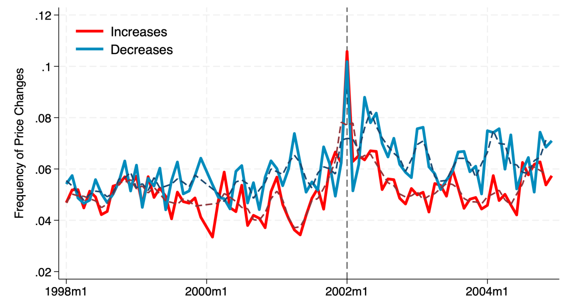
conversion from the national currency, the euro amount could be rounded to the nearest euro cent (two decimal places). For instance, the European Central Bank (ECB) set a fixed exchange rate for Croatia at 1 euro = 7.53450 kuna. Following this rate, HRK 100 would convert to EUR 13.2722808, which would then be rounded to EUR 13.27. These guidelines ensured consistency and minimized discrepancies during the transition. Since we are interested in studying price changes beyond those caused by rounding, we exclude changes smaller than 0.1%. This provides a common benchmark that can be applied consistently across all countries, including those for which we do not observe prices in the original currency. Because the Slovak data include price quotes in their original currency (Slovak koruna), we can exclude price changes attributable to rounding. Figure C5 shows that excluding price changes smaller than 0.1% is very similar to simply removing price changes caused by rounding. Similarly, as shown later using high-frequency data for Croatia, this restriction also eliminates most price changes resulting from rounding in that country.

We begin by presenting the patterns of pricing moments for Austria. Panel (a) in Figure 4 illustrates the frequency of price changes in Austria before and after the euro changeover, controlling for seasonality. The figure highlights a significant increase in the probability of price

Figure 4: Pricing Moments: Austria



(a) Frequency of Adjustment



(b) Frequency: Positive and Negative

(c) Absolute Size of Price Changes

(d) IQR

Note: The figure illustrates pricing moments for Austria before and after the euro changeover. Panel (a) shows the frequency of price changes after controlling for seasonality, with the dashed line representing the three-month moving average. Panel (b) depicts the frequency of price increases and decreases. Panel (c) reports the absolute size of price changes, while panel (d) presents the 25th and 75th percentiles of the price change distribution.

adjustments in the period leading up to the changeover.[15] The frequency of price changes increases for both price increases and price decreases, as shown in panel (b). The distribution of relative price changes also evolves as the changeover approaches. There is a significant reduction in the cross-sectional dispersion of price changes that begins a few months before the introduction of the euro and persists for a few months after the euro becomes the sole legal tender. Panel (c) shows that the absolute size of price changes decreases. Panel (d) depicts the first and third quartiles of the price-change distribution. The interquartile range, defined as the difference between the third and first quartiles, reaches its minimum at the time of the changeover and recovers in subsequent periods.[16] Figure C4 presents similar patterns observed in Slovakia, Greece, and Finland during the euro transition. For Austria, we have data on sales and product substitutions, while for Slovakia, the dataset includes information

[15] Figure C2 shows that these patterns are robust for price changes larger than 1% as well.

[16] Similar patterns can be seen in the full distribution of price changes shown in Figure C3.

on imputations and product substitutions. Figure C1 shows that the observed pricing patterns around the changeover are not influenced by sales, substitutions, or imputations, as these factors remain largely constant throughout the transition period.

Table 2: Pricing Moments at Changeover

Note: The table reports pricing moments before and during the changeover. For Austria, Finland, Greece, and the UK, the benchmark is the average across months at least seven months before the changeover within the 2000–2004 sample window, and the changeover period covers December 2001 and January 2002. For Slovakia, the benchmark is defined analogously within the 2007–2011 sample window, and the changeover period covers December 2008 and January 2009. To maintain comparability across countries, the baseline statistics include sales, since sales flags are available only for Austria, Finland, and the UK. Whenever the corresponding flags are available, price changes associated with substitutions, imputations, missing prices, or out-of-stock observations are not classified as price adjustments and do not enter the distribution of adjustment sizes. The statistics reported exclude price changes smaller than 0.1%. Frequency is the probability of price adjustment, excluding product substitution and imputations whenever such information is available. Size positive and size negative are the sizes of price change in logarithms. IQR is the difference between price changes in the 75th percentile and those in the 25th percentile.

Country	Period	(1) Freq.	(2) Freq. Pos.	(3) Freq. Neg.	(4) Size Pos.	(5) Size Neg.	(6) Abs. Size	(7) IQR
Austria	Benchmark	0.11	0.06	0.05	0.19	-0.24	0.21	0.31
	Changeover	0.17	0.09	0.08	0.19	-0.14	0.17	0.19
Finland	Benchmark	0.13	0.08	0.06	0.18	-0.21	0.19	0.20
	Changeover	0.22	0.12	0.10	0.19	-0.16	0.17	0.14
Greece	Benchmark	0.15	0.09	0.06	0.13	-0.14	0.14	0.16
	Changeover	0.27	0.16	0.11	0.12	-0.08	0.10	0.08
Slovakia	Benchmark	0.18	0.10	0.08	0.11	-0.13	0.12	0.17
	Changeover	0.32	0.16	0.15	0.10	-0.09	0.10	0.11
UK	Benchmark	0.13	0.07	0.06	0.16	-0.22	0.19	0.26
	Changeover	0.13	0.06	0.07	0.19	-0.23	0.21	0.30

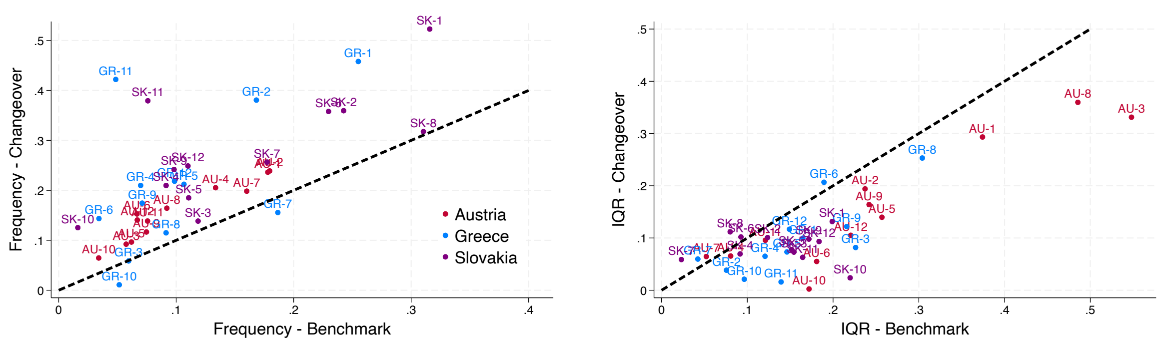
Table 2 summarizes these patterns and shows that they are similar across the countries transitioning to the euro. Column (1) reports the frequency of price adjustment during the December–January changeover period relative to a pre-changeover benchmark, defined as the average across months at least seven months before the changeover.[17] With the exception of the UK, which did not transition to the euro, the frequency of price changes is higher during the changeover. The frequency of adjustment rises from 0.11 to 0.17 in Austria, from 0.13 to 0.22 in Finland, from 0.15 to 0.28 in Greece, and from 0.18 to 0.32 in Slovakia. The increase in frequency is not confined to one direction: both price increases and price decreases become more frequent around the changeover. Columns (6) and (7) report two measures of the dispersion of price changes: the absolute size of log price changes and the interquartile

[17]For Austria, Finland, Greece, and the UK, the changeover period covers December 2001 and January 2002; for Slovakia, it covers December 2008 and January 2009.

range, computed as the difference between the 75th and 25th percentiles. In every adopting country, the average absolute size and interquartile range of price changes fall during the changeover. Table C1 shows that these patterns are robust to alternative definitions of price changes whenever the underlying data permit, including excluding changes attributable to mechanical rounding, salient prices, and sales.[18]

Sectors: The patterns in pricing moments that we uncover across countries also hold across sectors. Although the data sets differ in their sectoral aggregation, they provide information at the COICOP division level, corresponding to the 2-digit classification. Examples include Division 01 for "Food and Non-Alcoholic Beverages," Division 03 for "Clothing and Footwear," Division 07 for "Transport," Division 09 for "Recreation and Culture," and Division 10 for "Education."[19] Panel (a) in Figure 5 shows the frequency of price adjustment in each country and sector during the changeover relative to the corresponding benchmark, defined, as before, as the average across all months at least seven months before the changeover. The frequency of adjustment is higher during the changeover in nearly every country-sector cell, with the exceptions of Division 07 in Austria and Division 08 in Slovakia. Similarly, Panel (b) shows that the IQR of price changes is smaller during the changeover in most sectors and countries.

Figure 5: Pricing Moments by Sector



(a) Frequency (b) IQR

Note: Panel (a) shows the frequency of price adjustment for each COICOP division in Austria (AU), Greece (GR), and Slovakia (SK) during the changeover and the corresponding benchmark, defined as the average across months at least seven months before the changeover. Panel (b) uses the same periods and reports the interquartile range (IQR) of log price changes, defined as the difference between the 75th and 25th percentiles of the price-change distribution.

[18] Results are also robust to restricting the sample to a strictly balanced cohort of product-outlet pairs observed in every month from 13 months before through 12 months after the changeover.

[19] We exclude Finland from this analysis because access to the Finnish CPI microdata is restricted to on-site use at Statistics Finland, which prevents us from incorporating it into the country-sector dataset.

To systematically use the sectoral variation in our data, we next examine how pricing moments evolve around the introduction of the euro using an event-study approach. The euro changeover was not a random event, and its announcement may itself have affected pricing behavior before the conversion date. The event-study design allows us to trace these dynamics before and after the changeover while controlling for time-invariant observed and unobserved heterogeneity across country-sector cells. For this analysis, our sample period spans two years before and three years after the euro changeover across all countries, including Slovakia, whose changeover occurred at a different time compared to the other countries. The UK provides a historical non-changeover benchmark for all adopting countries. Thus, as a control group, we include sectors in the United Kingdom over the 2000–2004 period. Let Y_{ijt} represent an outcome variable for sector i, country j, and time t (e.g., frequency of price changes, interquartile range, etc.). The specification for our event study is as follows:

$$Y_{ijt} = \alpha + \sum_{\substack{k=-24 \\ k\neq -2}}^{35} \gamma_k \mathbb{1}\{j \in \mathcal{A}\} \mathbb{1}\{K_{ijt} = k\} + \sum_{m\neq 6} \delta_m \mathbf{1}\{m(t) = m\} + \theta_{ij} + \epsilon_{ijt} \tag{18}$$

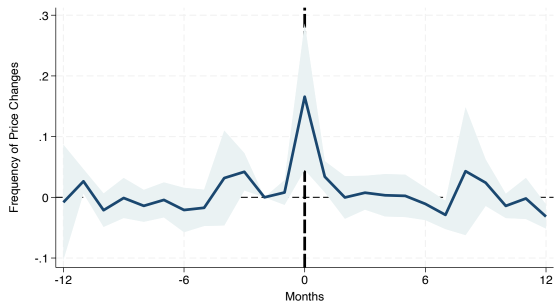
where $\mathcal{A} = \{\text{Austria, Greece, Slovakia}\}$ and θ_{ij} represents sector $\times$ country fixed effects. We considered two-digit COICOP codes as the relevant sectors. K_{ijt} denotes the number of months relative to the introduction of the euro, such that γ_k for $k < 0$ corresponds to pre-trends and $k \geq 0$ corresponds to dynamic effects k periods after the introduction of the euro. Event-time indicators are interacted with euro adoption and therefore equal zero for the UK. We control for seasonality using month-of-year effects, δ_m, and double-cluster the standard errors at both the country and sector levels.

Figure 6 shows the evolution of the main pricing moments around the introduction of the euro.[20] All estimates exclude small price changes and control for country-sector fixed effects and seasonality. Panel (a) shows little evidence of pre-trends in adjustment frequency, followed by a marked increase in the changeover month.[21] Panels (c) and (d) show that both

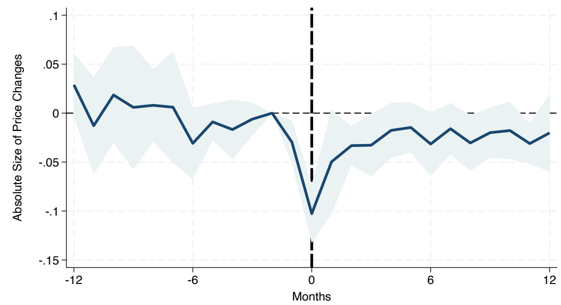
[20] Dual-pricing arrangements differed across countries, and the historical CPI files do not report both displayed prices. In Croatia, where we collected detailed information on dual-pricing practices, prices were closely monitored and any change linked to the currency conversion was prohibited. Not surprisingly, we see no unusual pricing activity during the dual-pricing period. More generally, its introduction does not generate noticeable movements in pricing moments in any of the countries we study. The European Commission left dual pricing to Member States, so practices varied widely. In many early adopters, including Finland, Germany, and France, dual pricing was voluntary and enforcement was generally lenient. We therefore describe these arrangements in Appendix D and align countries by the cash-conversion date, which is observed consistently across episodes.

[21] Because all changeovers in our sample occurred in January, month-of-year controls alone cannot distinguish the effect of the euro changeover from pricing behavior that is systematically unusual in January. In Appendix C.3, we address this concern by complementing the event-study analysis with placebo-January

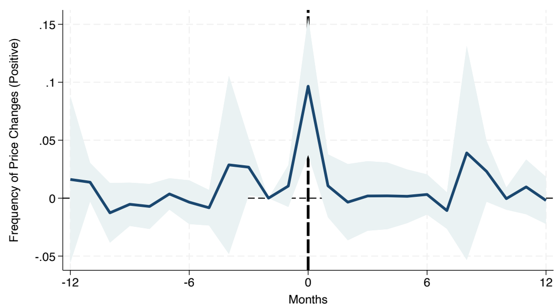
Figure 6: Event Study: Pricing Moments



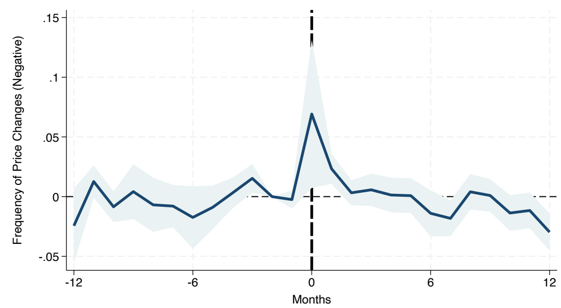
(a) Frequency



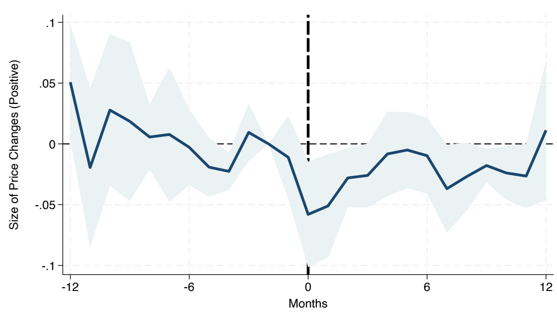
(b) Absolute Size



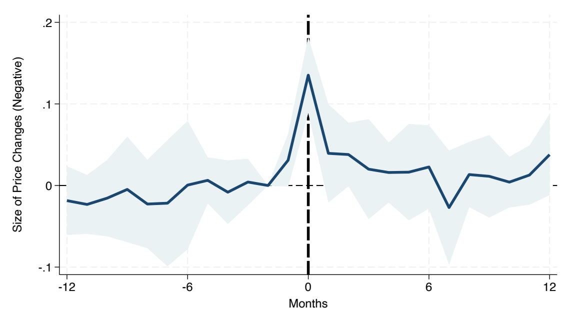
(c) Frequency Positive



(d) Frequency Negative



(e) Size Positive



(f) Size Negative

Note: The graph shows the evolution of several pricing moments before and after the introduction of the euro. The panels plot the coefficients of γ_k after estimating equation (18). The vertical dashed line indicates the month when the euro was introduced. Panel (a) depicts the frequency of price adjustments. Panel (b) shows the absolute size of price changes. Panels (c) and (d) present the frequency of price increases and decreases, respectively. Panels (e) and (f) display the size of price increases and decreases, respectively. The gray area represents the 95% confidence interval, calculated by clustering the standard errors at the country and sector levels.

price increases and price decreases contribute to this rise. Panel (b) shows a significant decline in the absolute size of price changes at the time of the conversion, while Panels (e) and (f)

randomization inference. We find that the actual increase in the frequency of price changes and decline in their absolute size are more extreme than those obtained under any of the placebo combinations.

show that both upward and downward adjustments become smaller. Replacing the month-of-year controls with sector-by-time fixed effects, which absorb shocks common to a sector across countries in each month, leaves the results essentially unchanged. Appendix Figure C6 shows the same sharp increase in adjustment frequency and decline in the absolute size of price changes at the changeover. These patterns are consistent with the model's prediction that the changeover raises the frequency of adjustment while compressing the size of price changes.

4.2 Recent Changeovers: Croatia and Bulgaria

Next, we study the most recent euro transitions, which took place in Croatia on January 1, 2023, and in Bulgaria on January 1, 2026. For both countries, we collect daily posted prices before and after the euro became the sole legal tender. In principle, studying high-frequency data offers important advantages relative to CPI micro data. First, they allow us to observe the timing of price adjustments much more precisely around the changeover date. Second, in the Croatian case, scraped price tags contain detailed information on how euro and kuna prices were displayed before and after the dual-pricing period.

However, the study of recent changeovers comes with important limitations. In particular, pricing during the transition was subject to unusually close monitoring as part of consumer-protection efforts. This is especially relevant in Bulgaria, where the data available to us are themselves reported by large retailers to the Consumer Protection Commission. The Commission monitored prices during the transition, and retailers could be required to justify price increases and faced sanctions for violations. Bulgaria also required large retailers to report and publish final sale prices daily.[22] As a result, the Bulgarian episode captures pricing behavior under an unusually intensive monitoring regime, making it less suitable for studying firms' unconstrained response to the changeover.

Croatia: We collect daily posted prices for 125,108 products (item-outlet) sold by 9 major retailers in Croatia. The data were collected through web scraping and include product identifiers, prices, and product categories. The categories cover a broad range of consumer goods, including recreation and culture, health, food and non-alcoholic beverages, and furnishing, household equipment, and routine maintenance. A pilot phase ran from January 2022 to May 2022, and the final data span May 2022 to July 2024. Importantly, the web-scraped prices closely match prices in the physical stores of the same retailers, supporting the reliability

[22] For Croatia, see https://www.ecb.europa.eu/euro/changeover/croatia/html/index.en.html. For Bulgaria, see Reuters, "Bulgaria sets price controls as euro transition nears," https://www.reuters.com/en/bulgaria-sets-price-controls-euro-transition-nears-2025-07-30/.

of the data used in the analysis.[23] In Croatia, compulsory dual pricing began in September 2022 to help consumers understand the new pricing system and verify the accuracy of conversions.[24] Retailers were required to display the exchange rate clearly and visibly alongside prices in both kunas and euros. There was a gradual shift in the prominence of the two currencies, first emphasizing the kuna price and later making the euro price more prominent after the changeover.[25]

Bulgaria: We collect daily posted prices from "Kolko Struva," a portal maintained by Bulgaria's Commission for Consumer Protection as part of the euro changeover. The portal publishes daily open-data files containing prices reported by large retail chains. The reporting requirement covers retailers with annual turnover above BGN 10 million, approximately EUR 5.1 million, selling products included in the large consumer basket defined by the Commission.[26] Importantly for our analysis, the data report both regular retail prices and promotional prices, allowing us to distinguish changes in regular prices from changes in the prices actually faced by consumers when promotions are available. The data cover the period from October 16, 2025, to June 8, 2026, and include 222 retailers, 6,481 stores, and 312,582 retailer-product identifiers.

Panel (a) in Figure 7 shows the frequency of price adjustments in Croatia, separately for increases and decreases, excluding changes smaller than 0.1%. The figure shows a clear concentration of price adjustments around the euro changeover date, consistent with a bunching of price changes in the last week before the changeover.[27] The increase in price decreases is

[23]To validate the web-scraped prices, we compared them with prices collected in physical stores in December 2022 and January 2023. Freelancers hired through Upwork visited stores and recorded prices for thirty items using the Mobile Inventory PRO free app; each freelancer received $20 per store visit. Immediately after receiving the offline data, we searched for the same products online to compare prices, minimizing the scope for price changes between the two measurements. Prices in physical stores closely matched those online. Before and after the euro changeover, both physical stores and online listings displayed prices in euros and kunas. Physical stores were more likely to display the two prices in the same size, while online listings typically made the euro price larger after the changeover. Most products observed offline were also available online, with an average match rate above 75%, stable before and after the changeover. When a product was available in both places, prices were almost always identical.

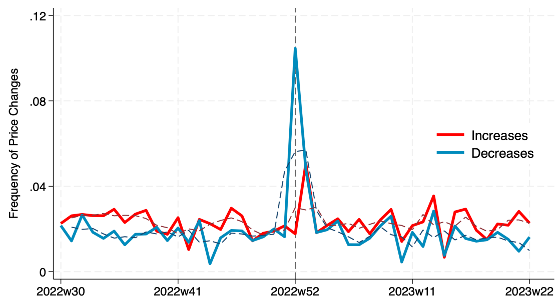
[24]The conversion from Croatian kuna (HRK) to euro (EUR) used the fixed exchange rate set by the European Central Bank: 1 euro = 7.53450 kuna. Prices were then rounded to the nearest cent. For example, HRK 15.00 converts to EUR 1.99084 and rounds to EUR 1.99.

[25]An example is shown in Figure C7, which presents an online listing from Decathlon in Croatia for the same item on December 30, 2022, and January 2, 2023, after the euro transition. The figure highlights a price change that occurred after the changeover.

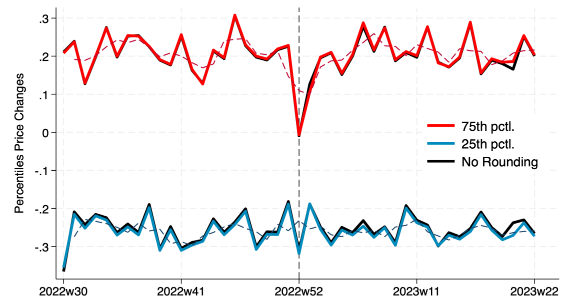
[26]Bulgaria introduced the euro on January 1, 2026, at a fixed conversion rate of 1 euro = 1.95583 Bulgarian lev. For details on the data, see `https://kolkostruva.bg/about` and `https://kolkostruva.bg/opendata`.

[27]Figure C8 shows the distribution of salient prices before and after the euro changeover for Croatia and Slovakia, two countries in our sample for which prices in the original national currency are observed prior to the transition. Salient prices, those with unusually common decimal endings such as 0, 5, or 9, stand out

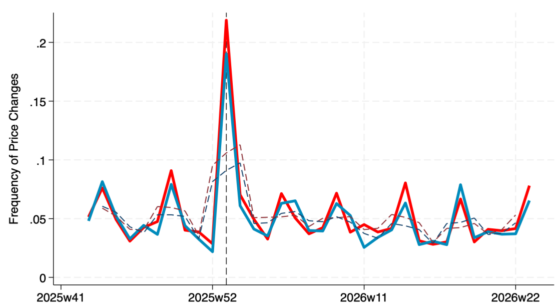
Figure 7: Pricing Moments: Croatia and Bulgaria



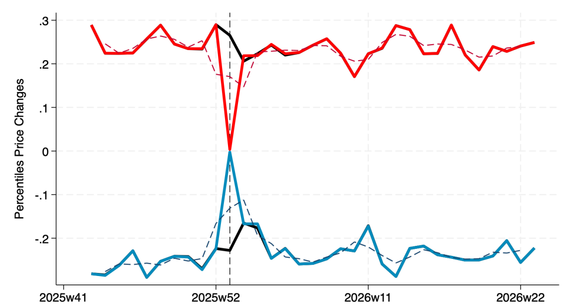
(a) Frequency: Croatia



(b) Distribution: Croatia



(c) Frequency: Bulgaria



(d) Distribution: Bulgaria

Note: The figure shows pricing movements in Croatia and Bulgaria before and after the euro changeover. Panels (a) and (c) present the frequency of price increases and decreases, excluding price changes smaller than 0.1%. Panels (b) and (d) display the 25th and 75th percentiles of the price-change distribution. The black lines show the corresponding percentiles after excluding changes fully accounted for by conversion into euros and rounding to the nearest cent. All prices include promotional prices.

concentrated in the changeover week, whereas price increases rise by much less and peak in the following week. This asymmetry is consistent with the consumer-protection rules in place during the transition, which placed particular scrutiny on price increases associated with the changeover.

Panel (c) shows the frequency of price changes in Bulgaria, including promotions. Using the common 0.1% threshold, there is a sharp spike in the first euro week. The spike remains when sales are excluded, suggesting that it is not driven by promotional activity. In this case, however, it nearly disappears once we exclude changes fully accounted for by conversion from lev to euros and rounding to the nearest cent.[28] The Bulgarian data are also unusual because

in the cross-sectional distribution of posted prices. At the changeover, the distribution of decimal endings becomes markedly flatter, and salient prices become substantially less common. Table C1 shows that the increase in adjustment frequency is robust to excluding changes to salient prices.

[28]See Table C1 for the corresponding moments under alternative definitions of a price change.

promotions account for a very large share of price changes. In normal weeks, prices change about 11% of the time, while regular prices change only about 1%, implying that promotions account for roughly 90% of observed price changes. Regular prices are therefore substantially stickier than in the other countries we examine.[29]

Panels (b) and (d) report the first and third quartiles of the distribution of price changes in Croatia and Bulgaria, respectively. In Croatia, the distribution compresses around the changeover, with a particularly sharp decline in the upper tail. In Bulgaria, the raw distribution also compresses sharply in the first euro week. Here too, however, the pattern is largely mechanical. Under the 0.1% threshold, the average absolute size of price changes falls from 0.28 to 0.19, but it is nearly unchanged once changes due solely to conversion and rounding are removed. This muted response may partly reflect the unusually restrictive setting in which the changeover took place. Large retailers reported prices daily to a public portal, could be asked to justify price increases on the basis of objective economic factors, and faced close scrutiny from both consumers and the Consumer Protection Commission. Discretionary price adjustments around the conversion were therefore unusually visible and potentially costly.

Given the unusual institutional environment surrounding the recent changeovers, we focus the remainder of the analysis on the historical CPI episodes, which provide a cleaner setting for studying firms' response to the changeover. In these episodes, the euro changeover is not neutral with respect to pricing moments. Firms do more than mechanically redenominate posted prices: they use the transition to adjust relative prices. The fact that these changes occur around a purely nominal event points to adjustment frictions in firms' price-setting decisions.

5 Comparing the Model to the Data

In this section, we compare the model's predictions with the data. Given the regulatory constraints in recent changeovers, we focus on CPI micro data from Austria, Greece, and Slovakia. The empirical tests use country-sector variation in the average absolute size and frequency of price changes. For Austria and Slovakia, sectors are defined at the four-digit COICOP level; for Greece, we use two-digit COICOP categories, which is the level available in the Greek CPI data.

Let $\bar{x}_{s,t}$ denote the average absolute size of price changes in country-sector cell s at time t, excluding price changes smaller than 0.1%. Let $\bar{x}_s$ denote the corresponding average before

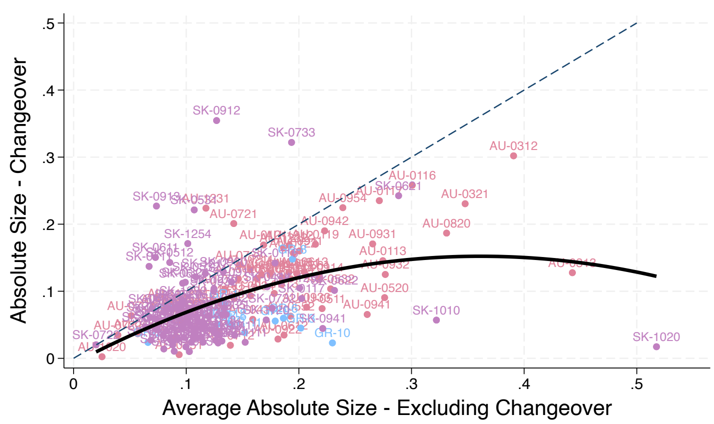
[29]For comparison, the weekly adjustment frequencies are 2.88% in Austria, 4.24% in Greece, 4.78% in Slovakia, 6.14% in Finland, and 6.29% in Croatia.

the changeover period, and let N_s denote the average frequency of price changes over the same benchmark period. The first test follows directly from Proposition 3. Since $\hat{\alpha}(T) = -3/2$, equation (14) implies that, just before the changeover,

$$\bar{x}_s(T) = \bar{x}_s - \frac{3}{2}\frac{\phi}{C}\frac{1}{\bar{x}_s} + o\left(\frac{\phi}{C}\right).$$

Thus, for a given value of ϕ/C, the size of price changes at the changeover should lie below its steady-state value, and the relationship between $\bar{x}_s(T)$ and $\bar{x}_s$ should be increasing and concave. Figure 8 plots this relationship across country-sector cells. The fitted quadratic lies below the 45-degree line and is concave over the relevant range. The corresponding quadratic regressions are reported in Appendix Table C2. In all specifications, the linear term is positive and the quadratic term is negative, and both are statistically significant. The estimates are similar with country fixed effects and when the constant is omitted. The estimated linear coefficient is approximately 0.85, and we cannot reject that it is equal to one, as predicted by the theory. Overall, the results support the model's prediction that adjustment sizes are compressed at the changeover and that the relationship is concave.

Figure 8: Absolute Size of Price Changes: Test of Proposition 3



Note: The figure plots the average absolute size of price changes during the changeover against the average absolute size of price changes before the changeover period (all months at least seven months before the changeover). Each observation is a country-sector cell. Sectors are defined at the four-digit COICOP level for Austria and Slovakia and at the two-digit COICOP level for Greece. The dashed line is the 45-degree line, and the solid line is a quadratic fit. Price changes smaller than 0.1% are excluded.

The second test uses the relative timing restriction in Proposition 3. For any $T_0 < t_1 < t_2 < T$, the ratio $\frac{\bar{x}_s(t_2)-\bar{x}_s}{\bar{x}_s(t_1)-\bar{x}_s}$ depends on the sector only through $N_s t_1$ and $N_s t_2$, and is otherwise independent of the remaining parameters. We take t_1 to be the month before the changeover and use the changeover month as the empirical counterpart of the limiting value as $t \uparrow T$.[30]

A direct implementation would form the ratio of deviations from the sectoral benchmark. In practice, this ratio is noisy because $\bar{x}_s(T-1) - \bar{x}_s$ is often close to zero, so small measurement errors generate very large values. We therefore implement the prediction as a level regression:

$$\bar{\chi}_{s,T} = \beta\, N_s \bar{\chi}_{s,T-1} + \mu_{j(s)} + \varepsilon_s,$$

where $\chi_{s,T} = \bar{x}_s(T) - \bar{x}_s$, $\chi_{s,T-1} = \bar{x}_s(T-1) - \bar{x}_s$, and $\mu_{j(s)}$ denotes a country fixed effect. This specification asks whether the changeover-month deviation is systematically related to the previous-month deviation scaled by the sector's normal frequency of adjustment, as suggested by the theory.

Table 3 reports the results. The estimated coefficient is positive in all specifications and is similar when country fixed effects are included and when the constant is omitted. Given the noise in sector-level deviations measured at a monthly frequency, we view this exercise as supporting evidence for the timing restriction in Proposition 3. In particular, sectors' normal frequency of adjustment helps account for the relationship between the compression in adjustment size before and at the changeover.

We next turn to a complementary exercise: a direct comparison between the model-implied dynamics and the empirical event-study patterns. We discipline the model using a simple calibration. As discussed in Section 3 there are two moments that inform the model calibration: the frequency of price changes, N, and the absolute size of the price changes, $\bar{x}$, both measured in the steady state. Given the observed steady state values of $N = 1.63$ (per year) and $\bar{x} = 0.17$, using a curvature parameter governing the quadratic loss equal to $\eta = 5$, implies a menu cost $\psi = 0.028$ as a fraction of annual profits, and $\sigma = 0.21$ for the volatility of the quality shocks.[31] We set the announcement horizon to $T = 1$ year. Relative to the steady-state model, the changeover introduces only one additional parameter, ϕ. We choose ϕ to match the frequency of price adjustment at the changeover. Setting $\phi = 0.07$ reproduces the observed increase in frequency to approximately 0.30. The path of adjustment sizes is not targeted in the calibration.

Figure 9 compares the model-implied time series of adjustment size and frequency to their

[30] We do not estimate T_0 directly. In the model, T_0 is the date at which euro adoption first becomes optimal conditional on a price adjustment, not the date of the first observed switch. Among the CPI files used in this section, only the Slovak data record the denomination of each price quote. At Slovakia's average pre-changeover adjustment frequency, the approximation in Proposition 2 places T_0 only a few months before the changeover. Consistent with this prediction, most denomination changes in Slovakia are recorded between the final pre-changeover and first post-changeover observations. The monthly data therefore provide little information about variation in T_0 across sectors.

[31] In the limit when $\rho \to 0$, the steady-state adjustment frequency satisfies $N = \sigma^2/\bar{x}^2$, and the inaction boundary is $\bar{x} = [6\psi\sigma^2/C]^{1/4}$.

Table 3: Cross-Sectional Test of Proposition 3

Note: The dependent variable is the deviation of the average absolute size of price changes in the changeover month from its benchmark value, $|\bar{x}(T)| - |\bar{x}|$. The regressor is the previous-month deviation, $|\bar{x}(T-1)| - |\bar{x}|$, interacted with the sector's steady state frequency of price adjustment, N. Columns (2) and (4) include country fixed effects. Columns (3) and (4) are estimated without a constant. The sample includes country-sector cells for Austria, Greece, and Slovakia. Sectors are defined at the four-digit COICOP level for Austria and Slovakia and at the two-digit COICOP level for Greece. Price changes smaller than 0.1% are excluded. Standard errors are reported in parentheses. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively.

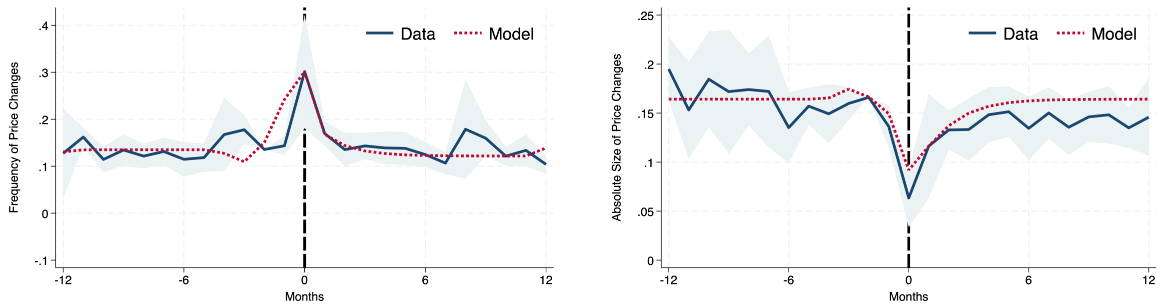
	(1)	(2)	(3)	(4)
		$\lvert\bar{x}(T)\rvert - \lvert\bar{x}\rvert$		
$N \times (\lvert\bar{x}(T-1)\rvert - \lvert\bar{x}\rvert)$	1.2344**	1.0647*	1.3957*	1.0647*
	(0.596)	(0.609)	(0.750)	(0.609)
Constant	-0.0578***	-0.0578***		
	(0.006)	(0.006)		
Observations	158	158	158	158
R-squared	0.027	0.053	0.022	0.053
Country	N	Y	N	Y
Constant	Y	Y	N	N

empirical counterparts around the changeover date. Panel (a) plots the frequency of price adjustments. The model closely matches the pronounced spike in adjustment frequency at $t = T$. After the changeover, both the model and the data exhibit a rapid return toward the steady-state adjustment rate. Panel (b) plots the absolute size of price changes. The model reproduces the gradual decline in adjustment size before the changeover, the sharp drop at T, and the subsequent recovery. Thus, with only one additional parameter for the changeover, the model captures the joint dynamics of frequency and size. This pattern reflects the rapid contraction of the inaction region predicted by the theory. Together with the cross-sectional evidence presented above, these results show that the model accounts for the central features of the data around the euro changeover.

5.1 What the Changeover Reveals About Price Flexibility

The exercises above also clarify what the changeover allows us to learn about price adjustment. In ordinary micro price data, observed price changes combine movements in firms' desired prices with the rule governing when firms choose to adjust. The euro conversion provides unusual variation because the fixed conversion rate leaves desired relative prices unchanged, while redenomination changes the payoff to taking a pricing action. The increase in adjustment frequency together with the decline in adjustment size therefore points to a change in the adjustment rule itself. In the model, the inaction region contracts and firms

Figure 9: Model vs. Data Around the Euro Changeover



(a) Frequency (b) Absolute Size

Note: The figure compares the model's predicted dynamics with the data around the euro changeover. Panel (a) shows the frequency of price changes, and Panel (b) shows the average absolute size of price changes. The calibration targets the steady-state frequency and average absolute size of price changes and the adjustment frequency at conversion. The path of adjustment sizes around conversion is not targeted.

with smaller price gaps are drawn into adjustment.

The results also show that this response has a simple time scale. A sector's steady-state adjustment frequency governs how quickly it responds to the changeover: sectors that normally adjust more often concentrate their response closer to the conversion date, while slower-adjusting sectors begin earlier. Thus, a steady-state pricing moment that usually summarizes price duration also contains information about the speed of adjustment away from steady state.

This distinction has become especially relevant with the recent return of high inflation, when the fraction of prices changing rose markedly. In those episodes, however, stronger incentives to adjust come together with movements in firms' desired prices, making the source of the frequency response harder to isolate. The changeover separates these forces and provides evidence on the state-dependent adjustment margin in a setting where the nominal conversion itself does not move desired relative prices.

6 Conclusion

This paper studies price-setting behavior around the euro cash changeover, a scheduled nominal reform in which firms transitioned their posted prices from the legacy currency to the euro. Because the conversion rate was fixed, the changeover did not by itself alter firms' desired relative prices. It therefore separates two things that ordinarily move together: the incentive to take a pricing action and the reason to change a real price.

We develop a nonstationary menu-cost model in which firms anticipate the conversion date and can combine a price adjustment with adoption of the new currency. As the changeover approaches, the value of adjusting, the inaction boundaries, and the distribution of price gaps evolve. The anticipated date makes the firm's problem a singular control problem with a moving boundary, which we solve in closed form by perturbation. The solution delivers the entire transition—the path of the inaction region, the date at which early adoption begins, and the implied pricing moments—and shows that it is governed by a single observable, the steady-state frequency of price adjustment N. The lead time of early adoption is $8\log 3/(\pi^2 N)$, and the compression of the inaction region at the conversion depends on the primitives only through N and the size of the mismatch cost. The model predicts that adjustment frequency rises while the size of price changes falls, with most of the response concentrated close to the conversion date.

We test these predictions using CPI micro data from Austria, Finland, Greece, and Slovakia, with the United Kingdom as a benchmark, together with evidence from the more recent Croatian and Bulgarian changeovers. In the historical episodes and in Croatia, adjustment frequency rises around the introduction of euro cash while the average absolute size and dispersion of price changes fall. In the historical CPI data, the relationship between adjustment size at the changeover and its steady-state value has the increasing and concave shape predicted by the theory, and the compression in adjustment size is related to sectors' normal frequency of price adjustment. Finally, a calibration disciplined by steady-state pricing moments and the changeover spike in adjustment frequency reproduces the decline in adjustment size and the dynamics around the conversion date.

These results bear on a difficulty that is intrinsic to the price-setting literature. Menu-cost and time-dependent models calibrated to the same steady-state moments deliver nearly identical responses to small monetary shocks, and their first-order inflation dynamics coincide, so the aggregate time series on which the literature usually relies cannot separate them. The changeover can, because it is large, anticipated, and purely nominal: it moves the value of taking a pricing action while leaving desired relative prices untouched, and firms respond. That they do is direct evidence of state dependence in the timing of price adjustment.

The distinction matters beyond currency changeovers, and has become more salient with the return of high inflation, when the share of firms changing prices rises along with the price level. It also matters for other reforms that are announced in advance and change the payoff to acting without changing what firms would like to charge—pre-announced tax changes, redenominations, dollarizations. The framework developed here applies to any of them.

References

Alvarez, F., Gonzalez-Rozada, M., Neumeyer, A., Beraja, M., 2011. From Hyperinflation to Stable Prices: ArgentinaÕs evidence on menu cost models. manuscript. University of Chicago.

Alvarez, F., Le Bihan, H., Lippi, F., 2016a. The real effects of monetary shocks in sticky price models: a sufficient statistic approach. American Economic Review 106, 2817–2851.

Alvarez, F., Lippi, F., Passadore, J., 2016b. Are state and time dependent models really different? NBER Macroeconomics Annual 2016, Volume 31 .

Alvarez, F., Lippi, F., Souganidis, P., 2023. Price setting with strategic complementarities as a mean field game. Econometrica 91, 2005–2039.

Auclert, A., Rigato, R., Rognlie, M., Straub, L., 2024. New pricing models, same old phillips curves? The Quarterly Journal of Economics 139, 121–186.

Barro, R.J., 1972. A theory of monopolistic price adjustment. Review of Economic Studies 39, 17–26.

Baumgartner, J., Glatzer, E., Rumler, F., Stiglbauer, A., 2005. How frequently do consumer prices change in austria? evidence from micro cpi data .

Berardi, N., Eife, T., Gautier, E., 2013. Optimal price setting during a currency changeover: Theory and evidence from french restaurants. Applied Economics 46.

Blanco, A., Boar, C., Jones, C., Midrigan, V., 2024. Non-linear inflation dynamics in menu cost economies .

Bonomo, M., Carvalho, C., 2004. Endogenous time-dependent rules and inflation inertia. Journal of Money, Credit and Banking 36, 1015–41.

Caballero, R.J., 1989. Time Dependent Rules, Aggregate Stickiness And Information Externalities. Discussion Papers 198911. Columbia University.

Calvo, G.A., 1983. Staggered prices in a utility-maximizing framework. Journal of Monetary Economics 12, 383–398.

Cavallo, A., Lippi, F., Miyahara, K., 2024. Large shocks travel fast. American Economic Review: Insights 6, 558–574.

Dixit, A., 1991. Analytical approximations in models of hysteresis. Review of Economic Studies 58, 141–51.

Ercolani, M.G., Dutta, J., 2006. The euro-changeover and euro-inflation: Evidence from eurostat's hicp. DEPARTMENT OF ECONOMICS DISCUSSION PAPER-UNIVERSITY OF BIRMINGHAM 3.

Gaiotti, E., Lippi, F., 2004. Pricing behavior and the introduction of the euro: Evidence from a panel of restaurants in Italy. Giornale degli Economisti 63, 491–526.

Golosov, M., Lucas, R.E.J., 2007. Menu costs and phillips curves. Journal of Political Economy 115, 171–199.

Hobijn, B., Ravenna, F., Tambalotti, A., 2006. Menu costs at work: restaurant prices and the introduction of the euro. The Quarterly Journal of Economics 121, 1103–1131.

Hüfner, F., Koske, I., 2008. The euro changeover in the slovak republic: implications for inflation and interest rates .

Karadi, P., Reiff, A., 2019. Menu costs, aggregate fluctuations, and large shocks. American Economic Journal: Macroeconomics 11, 111–146.

Kurri, S., 2007. Price changes in finland: some evidence from micro cpi data .

Meriküll, J., Rõõm, T., 2015. One currency, one price? euro changeover-related inflation in estonia. JCMS: Journal of Common Market Studies 53, 822–839.

Reis, R., 2006. Inattentive producers. Review of Economic Studies 73, 793–821.

APPENDIX

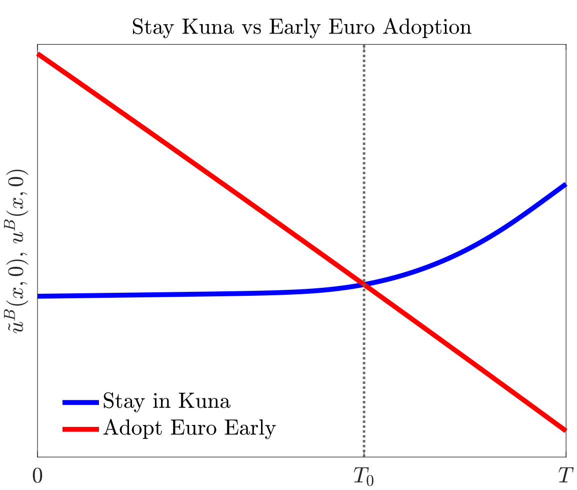
A Additional figures

This section contains additional figures describing the model's workings.

A.1 Value of Adjusting: kuna vs. euro

Figure A1 plots the value of adjusting prices at different times before the changeover, conditional on either staying in kuna or adopting the euro early. The horizontal axis shows calendar time t, and the vertical axis reports the value to the firm of adjusting at each t. The blue curve ("Stay in kuna") represents the minimum value function $u^B(0,t)$ conditional on remaining in kuna after adjustment. The red curve ("Adopt Euro Early") represents $\tilde{u}^B(0,t)$ conditional on adopting the euro. The value of early adoption declines away from T because adopting early requires paying the flow cost ϕ for longer, whereas the value of staying in kuna rises as the firm approaches the mandatory switch. The two curves intersect at T_0, marking the onset of early euro adoption.

Figure A1: Conditional Value of kuna vs euro for $t < T$

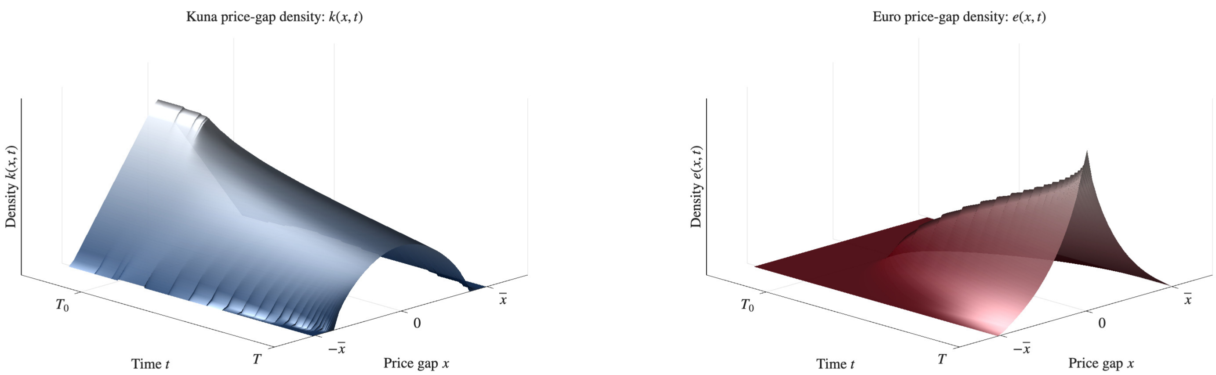


Note: The blue curve shows the value of an adjustment that keeps the firm in kuna ("Stay in kuna"), while the red curve shows the value of an adjustment that adopts the euro early ("Adopt euro early"). The curves intersect at the first adoption time T_0, marking the moment when euro adoption becomes optimal upon adjustment.

A.2 Evolution of the Cross-Sectional Distribution of Price Gaps

The figures below show the full time path of the cross-sectional densities $k(x,t)$ and $e(x,t)$ in three dimensions, complementing the two-dimensional slices in Figure 2 of the main text.

Figure A2: Evolution of Price-Gap Densities: Three-Dimensional View



Note: The left panel shows the density of price gaps $k(x,t)$ for firms still pricing in kuna, from T_0 to T. The right panel shows the density $e(x,t)$ for firms that have adopted the euro. As $t \to T$, the kuna density erodes from the boundaries as firms redenominate, while the euro density fans out from zero toward the stationary tent distribution. Both figures use the calibrated parameters reported in Section 5.

B Analytic Solution of the Perturbation

In this appendix we collect well known results for the solution of the one dimensional heat equation with given initial (or terminal) conditions, defined in a strip, with time varying boundaries, and allowing for a source.

Consider the initial value problem for the heat equation in the domain $(z,t) \in [0,1] \times \mathbb{R}_+$, with a source s, and with time boundaries given by the time varying functions A, B. In particular to solve for $w : [0,1] \times \mathbb{R}_+ \to \mathbb{R}$ given parameter $k > 0$, $\iota \geq 0$, source $s : [0,1] \times \mathbb{R}_+ \to \mathbb{R}$, space boundary at time zero $f : [0,1] \times \mathbb{R}$, and value at the boundaries given by $A, B : \mathbb{R}_+ \to \mathbb{R}$ satisfying:

$$0 = -w_t(z,t) - \iota w(z,t) + k w_{zz}(z,t) + s(z,t) \text{ all } z \in [0,1] \text{ and } t > 0 \tag{19}$$

$$w(z,0) = f(z) \text{ all } z \in [0,1] \tag{20}$$

$$w(0,t) = A(t) \text{ all } t > 0 \text{ and} \tag{21}$$

$$w(1,t) = B(t) \text{ all } t > 0 \tag{22}$$

Lemma 1. The solution for the KFE equation for w is given by:

$$w(z,t) = r(z,t) + \sum_{j=1}^{\infty} a_j(t)\varphi_j(z) \text{ all } z \in [0,1] \text{ and } t > 0 \text{ where} \tag{23}$$

$$r(z,t) = A(t) + z[B(t) - A(t)] \text{ all } z \in [0,1],\ t > 0 \tag{24}$$

and where for all $j = 1, 2, \ldots$ we have:

$$\varphi_j(z) = \sqrt{2}\sin(j\pi z) \text{ for all } z \in [0,1]\,,\ \langle \varphi_j, h\rangle \equiv \int_0^1 h(z)\varphi_j(z)dz \tag{25}$$

$$a_j(t) = a_j(0)e^{-\lambda_j t} + \int_0^t q_j(\tau)e^{\lambda_j(\tau - t)}d\tau \text{ all } t > 0\,, \tag{26}$$

$$q_j(t) = \langle \varphi_j\,,\ s(\cdot,t) - r_t(\cdot,t) - \iota r(\cdot,t)\rangle \text{ all } t > 0 \tag{27}$$

$$\lambda_j = \iota + (j\pi)^2 k \ \text{ and } a_j(0) = \langle \varphi_j\,,\ f - r(\cdot,0)\rangle\,. \tag{28}$$

The proof can be done by verifying that the equation holds at the boundaries, that for $t > 0$ the p.d.e. holds in the interior since

$$a_j'(t) = -\lambda_j a_j(t) + q_j(t) \text{ for all } t > 0 \text{ and } j = 1, 2, \ldots$$

and since $\{\varphi_j(z)\}$ form an orthonormal basis for functions on $\{h : [0,1] \to \mathbb{R}\}$, and finally that the boundary holds at $t = 0$ for all $z \in [0,1]$.

B.1 Proof of Proposition 1

We now derive the closed-form solution for $v^B(x,t)$ stated in Proposition 1. The starting point is the heat equation (9) with Dirichlet boundary conditions in equation (10).

Mapping to the general framework. There is no internal source term ($s = 0$). The terminal condition is $v^B(x,T) = v^A(x)$ at $t = T$, and the sign of the time derivative is reversed when mapping into an initial-value problem, since $t = 0$ in the general statement corresponds here to $t = T$ (so the original time index becomes $T - t$). We set $\iota = \rho$, and the Dirichlet boundary conditions $v^B(\pm\bar{x}, t)$ become $A(t) = B(t) = \frac{1 - e^{-\rho(T-t)}}{\rho}$, so that $r(z,t) = A(t)$ in the notation of Lemma 1. Note that $r_t(z,t) - \rho r(z,t) = -1$, which simplifies the coefficients q_j since the boundary conditions do not depend on z and $A(T) = 0$.

The state variable is $z \in (0,1)$, corresponding to the normalized original state variable

$x \in (-\bar{x}, \bar{x})$, defined by $z \equiv (\bar{x}+x)/(2\bar{x})$. Letting $w(z,t) = v^B(x,t)$ and applying equation (9) yields the PDE:

$$\begin{aligned}
\rho w(z,t) &= w_t(z,t) + k\, w_{zz}(z,t) \qquad \text{for all } z \in [0,1],\ t > 0, \qquad k = \frac{\sigma^2}{8\bar{x}^2}, \\
w(z,T) &= v^A(\bar{x}(2z-1)) \qquad \text{for all } z \in [0,1], \\
w(0,t) &= A(t), \qquad w(1,t) = A(t) \qquad \text{for all } t \in [0,T].
\end{aligned}$$

Intermediate solution. Substituting into Lemma 1 yields:[32]

COROLLARY 1. The solution for $w(z,t)$ is

$$w(z,t) = A(t) + \sum_{j=1}^{\infty} a_j(t)\, \varphi_j(z),$$

for all $z \in [0,1]$ and $t \in [0,T]$, where $\varphi_j(z) = \sqrt{2}\sin(j\pi z)$ and

$$\begin{aligned}
a_j(t) &= a_j(T)e^{-\lambda_j(T-t)} + \int_t^T q_j(\tau)e^{\lambda_j(t-\tau)}\, d\tau, \\
q_j(t) &= \langle \varphi_j,\, A_t(t) - \rho A(t)\rangle = \langle \varphi_j, -1\rangle = \begin{cases} 0 & \text{if } j \text{ even}, \\ -\frac{2\sqrt{2}}{j\pi} & \text{if } j \text{ odd}, \end{cases} \\
\lambda_j &= \rho + (j\pi)^2 \frac{\sigma^2}{8\bar{x}^2}, \qquad a_j(T) = \langle \varphi_j, v^A\rangle.
\end{aligned}$$

Projection coefficients. We compute $a_j(T) = \langle \varphi_j, v^A\rangle$:

$$\begin{aligned}
a_j(T) &= \frac{\sqrt{2}}{\rho}\left(\frac{1-\cos(j\pi)}{j\pi} - \frac{2\big(j\pi\, \sin^2(\frac{j\pi}{2})\big)}{4(\xi\bar{x})^2 + (j\pi)^2}\right) \\
&= \begin{cases} 0 & \text{if } j \text{ even}, \\ \frac{2\sqrt{2}}{j\pi\rho}\left(1 - \frac{(j\pi)^2}{4(\xi\bar{x})^2+(j\pi)^2}\right) & \text{if } j \text{ odd}. \end{cases}
\end{aligned}$$

[32] The solution can be verified by checking that it satisfies the boundary conditions and that for $t \in [0,T]$ the PDE holds in the interior because $-a_j'(t) = -\lambda_j a_j(t) + q_j(t)$ for all $t \in [0,T)$, $j = 1, 2, \ldots$. Note that $q_j(t)$ and $a_j(t)$ differ from those in Lemma 1 because of the change of variable $t' = T - t$.

Also note that

$$\int_t^T q_j(\tau) e^{\lambda_j(t-\tau)}\, d\tau = \begin{cases} 0 & \text{if } j \text{ even,} \\ -\frac{2\sqrt{2}}{j\pi} \frac{1-e^{\lambda_j(t-T)}}{\lambda_j} & \text{if } j \text{ odd.} \end{cases}$$

Combining these expressions yields Proposition 1. The optimal threshold $\bar{x}(t)$ in equation (12) then follows from the perturbation of the smooth-pasting condition $u^B_{xx}(\bar{x}, t, 0)\,\alpha(t) + v^B_x(\bar{x}, t) = 0$ evaluated using $\partial v^B/\partial x = (\partial w/\partial z) \cdot 1/(2\bar{x})$.

Proof of approximate T_0. For part *(ii)*, use the T_0 definition, $\tilde{u}^B(0, T_0) = u^B(0, T_0)$, the expressions for the value function $u^B(0, T_0)$, the one for the approximate value function $u^B(0, T_0) = \tilde{u}^A(0, T_0) + \phi v^B(0, T_0)$, and the solution for V^B given above, to obtain that T_0 satisfies $\sum_{j=1,3,5,\ldots} b_j(T_0)\, \sin\left(\frac{j\pi}{2}\right) = 0$.

In the limit $\rho \to 0$, substituting the limiting expression for b_j:

$$\sum_{j=1,3,5,\ldots} \frac{4}{j\pi} \cdot \frac{3e^{-\lambda_j(T-T_0)} - 1}{\lambda_j} \cdot \sin\left(\frac{j\pi}{2}\right) = 0.$$

Since $\sin(j\pi/2)$ alternates as $+1, -1, +1, -1, \ldots$ for $j = 1, 3, 5, 7, \ldots$ and the coefficients $\frac{4}{j\pi} \cdot \frac{1}{\lambda_j} \sim j^{-3}$ decay rapidly, the sum is dominated by the $j = 1$ term. Setting it to zero gives

$$3e^{-\lambda_1(T-T_0)} - 1 = 0 \quad \Longrightarrow \quad T - T_0 = \frac{\log 3}{\lambda_1} = \frac{8 \log 3}{\pi^2 N},$$

where we used $\lambda_1 = \pi^2 N/8$ in the last step.

■

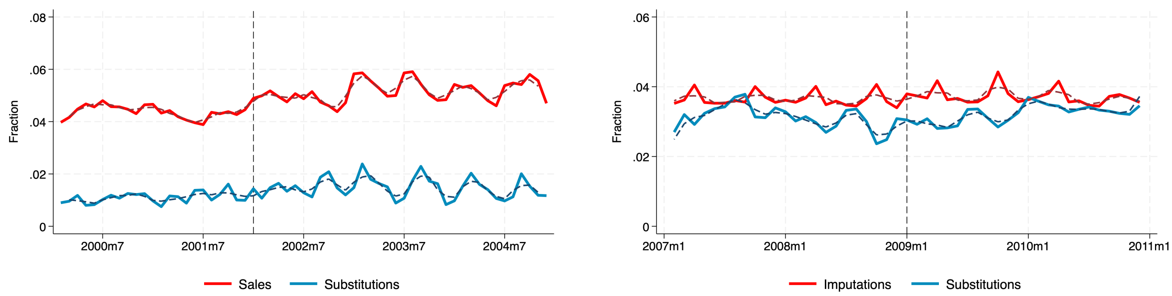
C Empirical Appendix

C.1 Figures

Figure C1: Sales, Imputations, and Substitutions: Austria and Slovakia

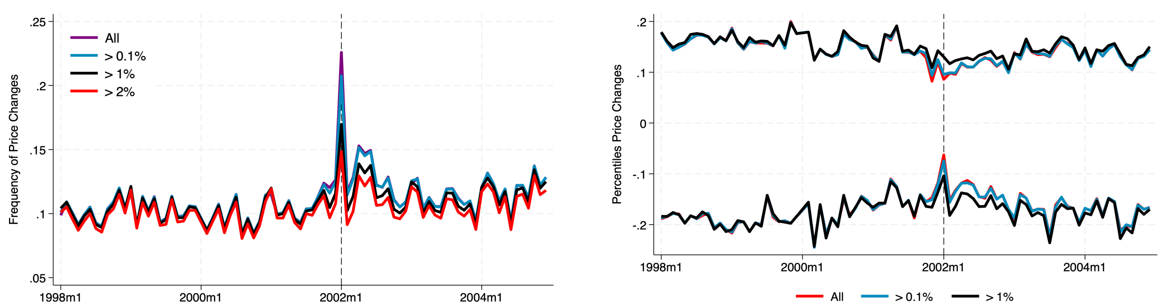
(a) Austria

(b) Slovakia



Note: Panel (a) shows the frequency of sales and product substitutions in Austria around the euro changeover. Panel (b) shows the frequency of price imputations and product substitutions in Slovakia around the euro changeover. These indicators are provided by each data set respectively. Observations are aggregated using equal weights for each item.

Figure C2: Pricing Moments: Austria

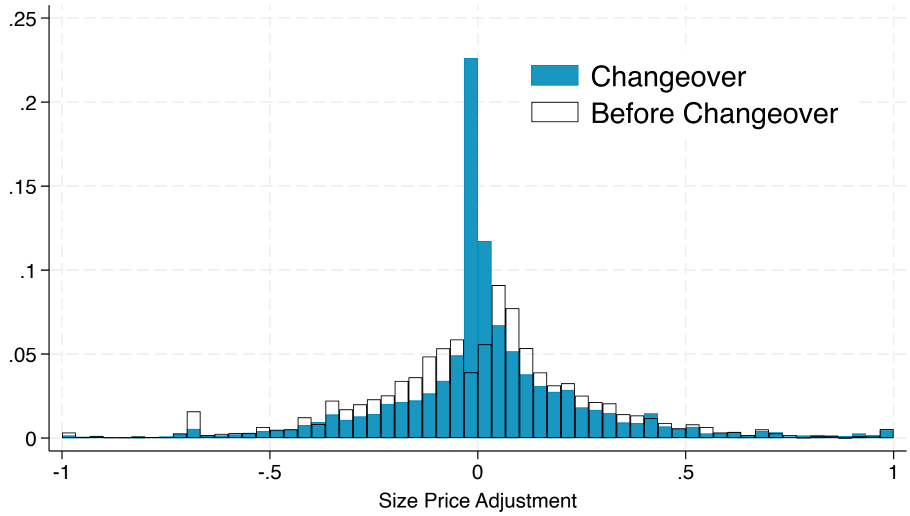


(a) Frequency of Adjustment

(b) IQR

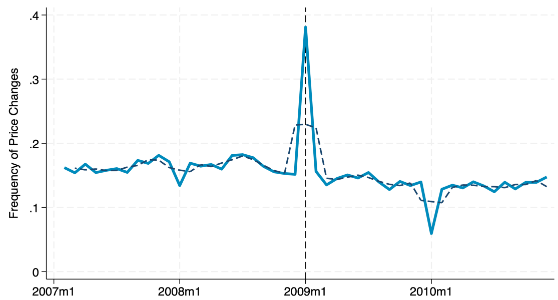
Note: The figure shows pricing moments for Austria before and after the euro changeover. Panel (a) presents the frequency of price changes after controlling for seasonality, with the dashed line representing the three-month moving average. Panel (b) shows the 25th and 75th percentiles of the price change distribution. The purple line depicts the frequency of all price changes, the blue line reports price adjustments larger than 0.1%, and the black line represents price changes larger than 1%.

Figure C3: Distribution of Price Changes: Austria



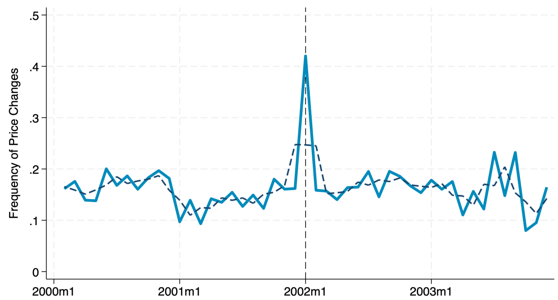
Note: The figure shows the distribution of price changes in December–January of 2002 and in the same months one year earlier. The histogram includes all price changes, including sales.

Figure C4: Pricing Moments: Slovakia, Greece, and Finland

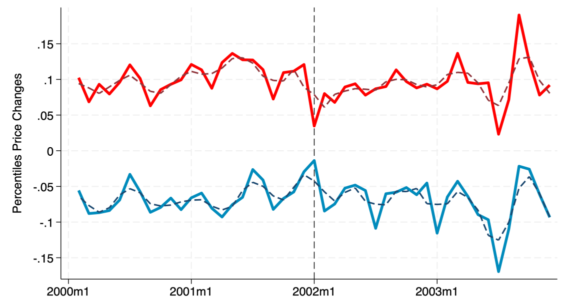


(a) Frequency: Slovakia

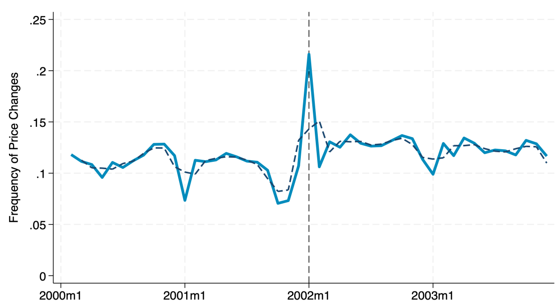
(b) IQR: Slovakia



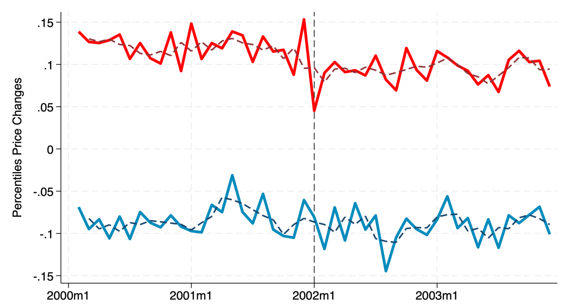
(c) Frequency: Greece



(d) IQR: Greece



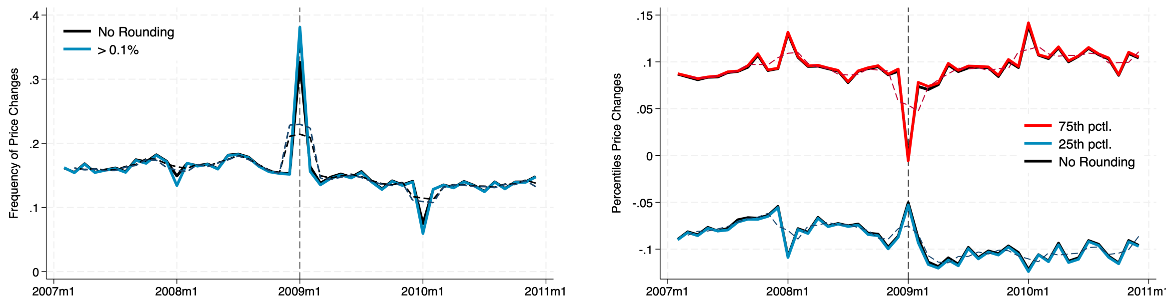
(e) Frequency: Finland



(f) IQR: Finland

Note: The figure shows pricing moments for Slovakia, Greece, and Finland before and after the euro changeover. Panels (a), (c) and (e) present the frequency of price changes after controlling for seasonality, with the dashed line representing the three-month moving average. Panels (b), (d) and (f) display the 25th and 75th percentiles of the price change distribution.

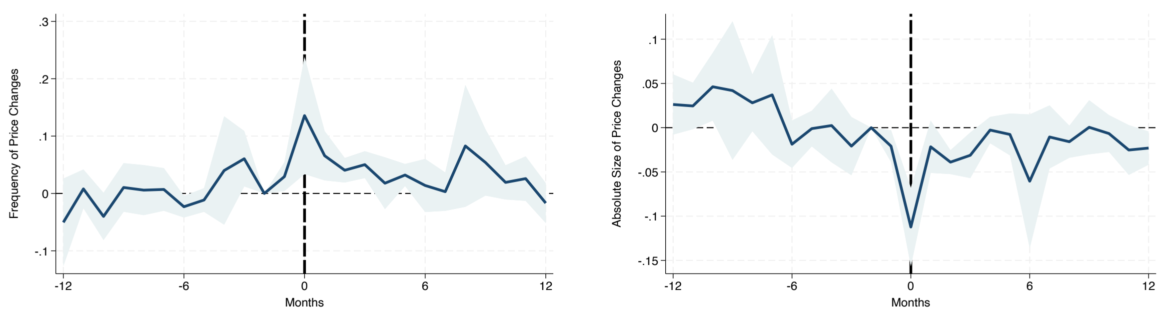
Figure C5: Pricing Moments: Slovakia (No Rounding)



(a) Frequency of Adjustment (b) IQR

Note: The figure shows pricing trends for Slovakia before and after the euro changeover. Panel (a) presents the frequency of price changes as a three-month moving average after controlling for seasonality. Panel (b) displays the 25th and 75th percentiles of the price change distribution. The red line depicts the frequency of all price changes excluding those due to rounding, while the blue line represents price adjustments larger than 0.1%.

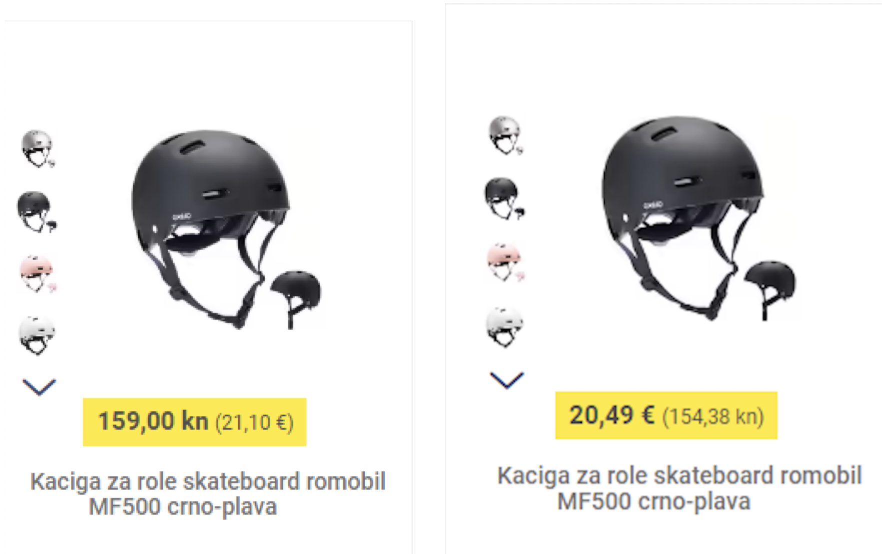
Figure C6: Event Study: Sector $\times$ Time Fixed Effects



(a) Frequency of Adjustment (b) Absolute Size of Price Changes

Note: The figure reports event-study estimates from a specification that includes country-sector and sector-by-time fixed effects. The secto $\times$ time fixed effects are defined at the two-digit COICOP-by-calendar-month level. Panel (a) reports the frequency of price changes, and panel (b) reports their average absolute size. The control group includes UK CPI data around both changeover dates: the January 2002 changeovers in Austria and Greece, and Slovakia's January 2009 changeover. Price changes smaller than 0.1% are excluded. The shaded areas represent 95% confidence intervals, with standard errors clustered by country and sector. The vertical dashed line marks the changeover month.

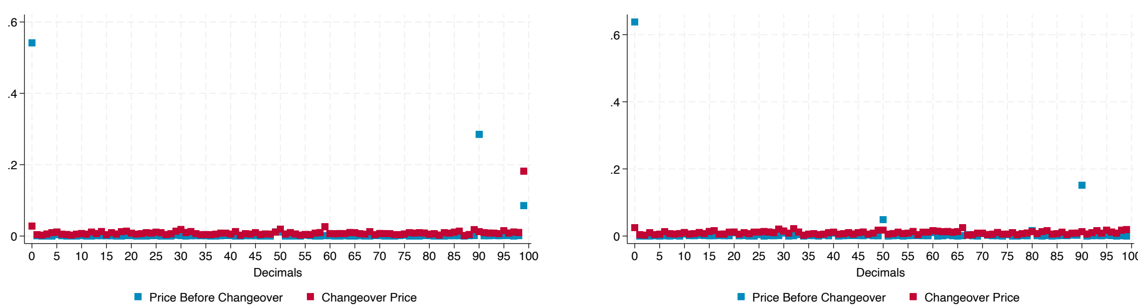
Figure C7: Croatia: Example of a Listing



(a) Dec 30, 2022 (b) Jan 2, 2023

Note: The figure shows an example of an online listing for a roller skate helmet at Decathlon in Croatia on December 30, 2022, and January 2, 2023, i.e., before and after the euro transition.

Figure C8: Salient Prices: Croatia and Slovakia



(a) Croatia (b) Slovakia

Note: The histograms plot the distribution of decimal endings for posted prices before the changeover (blue) and at the changeover (red).

C.2 Tables

Table C1: Pricing Moments at Changeover: Alternative Definitions

Note: Panel A uses historical monthly CPI data. For Austria, Finland, and Greece, the benchmark is the average across months at least seven months before the changeover within the 2000–2004 sample window, and the changeover period covers December 2001–January 2002. For Slovakia, the benchmark is defined analogously within the 2007–2011 sample window, and the changeover period covers December 2008–January 2009. Panel B uses weekly data. For Croatia, the benchmark covers 2022w18–2022w49 and the changeover period covers 2022w52–2023w1. For Bulgaria, the benchmark covers 2025w42–2025w50 and the changeover period covers 2025w52–2026w1. No rounding excludes price changes that are fully accounted for by conversion into euros and rounding to the nearest cent. Salient prices are prices ending in 00, 50, 90, or 99 cents. No sales specification excludes promotional prices.

		Frequency			Absolute Size		
Country	Definition	Benchmark	Changeover	Change	Benchmark	Changeover	Change
Panel A: Monthly CPI data							
Austria	0.1 percent	0.11	0.17	0.06	0.21	0.17	-0.05
	No sales	0.08	0.14	0.06	0.19	0.14	-0.05
	No salient	0.10	0.14	0.04	0.21	0.17	-0.04
Greece	0.1 percent	0.15	0.27	0.12	0.14	0.10	-0.03
	No salient	0.15	0.25	0.10	0.14	0.10	-0.03
Slovakia	0.1 percent	0.18	0.32	0.14	0.12	0.10	-0.02
	No rounding	0.18	0.28	0.10	0.12	0.10	-0.02
	No salient	0.17	0.29	0.12	0.12	0.09	-0.02
Finland	0.1 percent	0.13	0.22	0.09	0.19	0.17	-0.02
	No sales	0.09	0.18	0.09	0.17	0.15	-0.02
Panel B: High-frequency data							
Croatia	0.1 percent	0.06	0.20	0.14	0.24	0.07	-0.17
	No rounding	0.06	0.17	0.10	0.24	0.09	-0.15
	No salient	0.01	0.10	0.09	0.21	0.06	-0.15
Bulgaria	0.1 percent	0.11	0.27	0.16	0.28	0.19	-0.09
	No rounding	0.11	0.12	0.01	0.28	0.27	-0.01
	No salient	0.08	0.25	0.17	0.27	0.18	-0.09
	No sales	0.01	0.19	0.17	0.12	0.05	-0.07

Table C2: Quadratic Test of the Cross-Sectional Prediction of Proposition 3

Note: The dependent variable is the average absolute size of log price changes in the changeover month, $|\bar{x}(T)|$. The regressors are the country-sector's benchmark average absolute size of price changes, $|\bar{x}|$, and its square. The benchmark is computed as the average across all months at least seven months before the changeover. Columns (2) and (4) include country fixed effects. Columns (3) and (4) are estimated without a common constant. The sample includes country-sector cells for Austria, Greece, and Slovakia. Sectors are defined at the four-digit COICOP level for Austria and Slovakia and at the two-digit COICOP level for Greece. Price changes smaller than 0.1% are excluded. Standard errors are reported in parentheses. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively.

	(1)	(2)	(3)	(4)
	$\lvert\bar{x}(T)\rvert$			
$\lvert\bar{x}\rvert$	0.8845***	0.8497***	0.8086***	0.8497***
	(0.192)	(0.200)	(0.063)	(0.200)
$\lvert\bar{x}\rvert^2$	-1.2242***	-1.2134***	-1.0743***	-1.2134***
	(0.430)	(0.436)	(0.240)	(0.436)
Constant	-0.0076	-0.0027		
	(0.018)	(0.019)		
Observations	158	158	158	158
R-squared	0.224	0.263	0.730	0.263
Country	N	Y	N	Y
Constant	Y	Y	N	N

C.3 Placebo-January Randomization Inference

Because the changeovers in our sample occurred in January, month-of-year controls alone cannot distinguish the effect of the euro changeover from pricing behavior that is systematically unusual in January. We therefore complement the event-study analysis with a placebo exercise that compares the actual changeovers with other Januaries observed in the data.

We conduct the exercise jointly across Austria, Greece, and Slovakia. For each country c and candidate January in year y, we construct

$$C_{cy} = \left(Y_{c,\mathrm{Jan}(y)} - Y_{UK,\mathrm{Jan}(y)}\right) - \left(Y_{c,\mathrm{Nov}(y-1)} - Y_{UK,\mathrm{Nov}(y-1)}\right).$$

Thus, each C_{cy} measures the change from November to January in the treated country relative to the corresponding change in the United Kingdom. This preserves the same calendar comparison and pre-period normalization used for the actual changeover. We then average these country-level changes across Austria, Greece, and Slovakia. To construct the placebo distribution, we assign each country one of its non-changeover Januaries, calculate the same statistic, and average across the three countries. Considering every possible assignment produces 90 placebo combinations. The two-sided randomization probability is

$$P_R = \frac{1 + \sum_{b=1}^{90} \mathbf{1}\left\{|S^{(b)}| \geq |S^{\text{obs}}|\right\}}{91}.$$

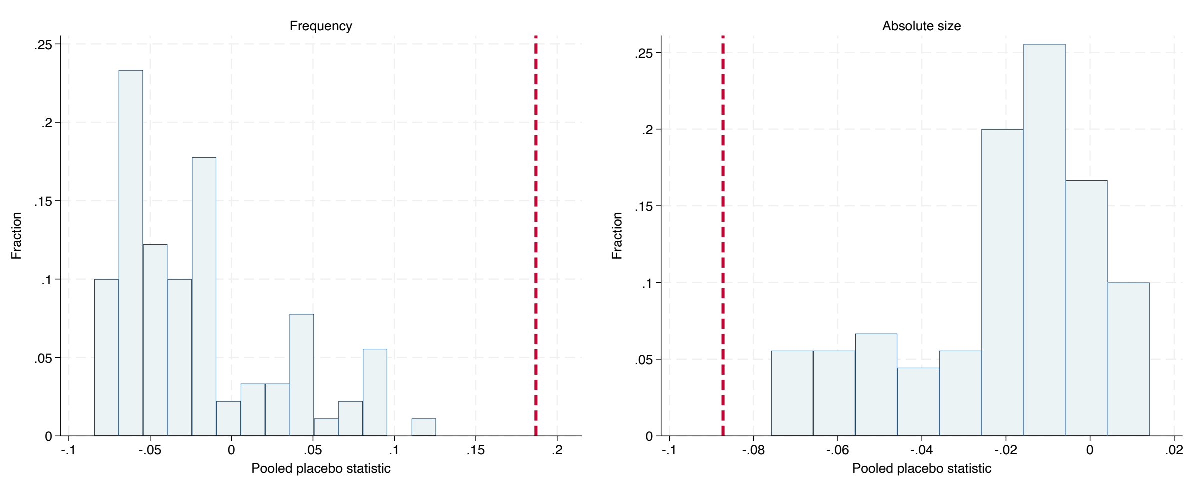
Table C3: Placebo-January Randomization Inference

Note: Each placebo assignment selects one non-changeover January for Austria, Greece, and Slovakia. For each country, we calculate the January-minus-November change relative to the corresponding change in the United Kingdom and then average across the three countries. The table reports the randomization p-value based on all 90 possible placebo assignments.

Outcome	Moment	Randomization p-value	Combinations
Frequency	0.187	0.011	90
Absolute size	-0.087	0.011	90

Table C3 and Figure C9 report the results. The increase in the frequency of price changes is 0.187, while the actual change in their absolute size is -0.087. Both statistics are more extreme than every one of the 90 placebo combinations, corresponding to a randomization p-value of 0.011 for both outcomes. The results indicate that the changes observed at the euro changeover are unusual relative to the pricing changes observed in other Januaries under the same empirical design.

Figure C9: Placebo-January Distributions



Note: The distributions contain all 90 placebo assignments. The dashed vertical line denotes the statistic for the euro changeovers. The increase in the frequency of price changes and the decline in their absolute size are more extreme than all placebo assignments.

D Introduction of the Euro: Timetable

The euro adoption process in Austria, Finland, and Greece differed from that in Slovakia, Croatia, and Bulgaria. Austria, Finland, and Greece followed the "Madrid scenario," with a transitional period during which the euro was the official currency but existed only as book money. The euro was introduced as the official currency on January 1, 1999 in Austria and Finland, and on January 1, 2001 in Greece. Euro banknotes and coins were then introduced in all three countries on January 1, 2002. The euro and the legacy currencies circulated in parallel through February 28, 2002. Dual-pricing arrangements differed across the three countries and are described separately below.

By contrast, Slovakia, Croatia, and Bulgaria followed a "Big Bang" approach, under which the euro became the official currency and euro banknotes and coins entered circulation on the same date, without a prior book-money transition. In each case, there was a short period of dual circulation and a longer period of compulsory dual pricing. Slovakia introduced the euro on January 1, 2009, Croatia on January 1, 2023, and Bulgaria on January 1, 2026.

Austria

- 1995: EU membership.
- 1995: ERM (European Exchange Rate Mechanism) membership.
- 1999: EMU (Economic and Monetary Union) membership.
- January 1, 1999: The euro was introduced as the official currency but only existed as "book money".
- September 1, 2001: Euro information campaign. Start of euro frontloading to credit institutions and possible start of euro sub-frontloading to companies.
- October 1, 2001: Dual pricing starts.
- December 15, 2001: Euro starter kits for consumers.
- January 1, 2002: Euro banknotes and coins are introduced in Austria.
- January 1, 2002: Dual circulation begins – Austrian schillings and euro are both legal tender.
- February 28, 2002: Dual circulation ends – the euro is the sole legal tender; exchange of schillings into euro is only possible at the OeNB (Oesterreichische Nationalbank) without amount and time limits.

- February 28, 2002: End of dual pricing.

Finland

- 1995: EU membership.
- 1996: ERM (European Exchange Rate Mechanism) membership.
- December 1997: The final changeover plan is published.
- 1999: EMU (Economic and Monetary Union) membership.
- January 1, 1999: The euro is introduced as the official currency but only exists as "book money".
- December 1999: The Bank of Finland initiates information efforts for the cash changeover and starts regular cooperation with the national campaign.
- October 1, 2001: Voluntary dual pricing starts.
- January 1, 2002: Euro banknotes and coins are introduced in Finland.
- January 1, 2002: Dual circulation begins.
- February 28, 2002: Dual circulation ends – the euro is the sole legal tender.
- February 28, 2002: End of dual pricing.

Greece

- 1981: EU membership.
- January 1, 1999 – December 31, 2000: ERM II membership.
- January 2001: EMU (Economic and Monetary Union) membership.
- January 1, 2001: The euro is introduced as the official currency but only exists as "book money". The adoption of the single currency marks the start of the "transitional period", in which the euro can only be used in scriptural form and the domestic market forms part of the integrated euro area market. During the transitional period, monetary policy is conducted in euro, credit institutions' current accounts with the Bank of Greece are denominated in euro, and all transactions in the interbank money market are carried out in euro. The transitional period lasts one year and ends on December 31, 2001.

- September 1, 2001: The Bank of Greece supplies commercial banks with euro coins.
- October 1, 2001: The Bank of Greece supplies commercial banks with euro banknotes. Dual pricing starts.
- November 1, 2001: Banks that have been frontloaded with euro coins and banknotes by the Bank of Greece sub-frontload euro coins to retailers.
- December 1, 2001: Banks that have been frontloaded with euro coins and banknotes by the Bank of Greece sub-frontload low-denomination banknotes (€5 and €10) to retailers.
- December 17, 2001: The National Cash Changeover Plan provides for the frontloading to the public of starter kits containing 45 euro coins of all denominations worth 5,000 drachmas (€14.67).
- January 1, 2002: Euro banknotes and coins are introduced in Greece.
- January 1, 2002: Dual circulation begins – euro banknotes and coins circulate in parallel with drachma banknotes and coins. Non-cash transactions (cheques, remittances, credit cards, etc.) are conducted in euro only.
- February 28, 2002: Dual circulation ends – the euro is the sole legal tender.

Slovakia

- May 2004: EU membership.
- 2005: ERM II membership.
- July 2008: Conversion foreign exchange rate of SKK/EUR set by the EU Council.
- July–December 2008: Ensuring the necessary number of euro banknotes and coins for cash circulation in Slovakia.
- September–December 2008: Supplying the National Bank of Slovakia and commercial banks with cash euros.
- December 2008: Supplying retailers with cash euros.
- August 24, 2008: Dual pricing starts.
- December 31, 2009: End of dual pricing.

- January 1, 2009: Accession into the euro area and introduction of the euro.
- January 16, 2009: Dual circulation ends (exchange of crown cash for euros and withdrawal of Slovak crowns from circulation). The euro becomes the sole legal tender.
- June 2010: End of recommended dual pricing.

Croatia

- July 1, 2013: EU membership.
- July 10, 2020: ERM II (European Exchange Rate Mechanism II) membership.
- July 2021: Conversion foreign exchange rate of HRK/EUR set by the EU Council.
- 2021–2022: Ensuring the necessary number of euro banknotes and coins for cash circulation in Croatia.
- October 2022–December 2022: Supplying the Croatian National Bank and commercial banks with cash euros.
- December 2022: Supplying retailers with cash euros.
- September 5, 2022: Dual pricing starts.
- January 1, 2023: Accession into the euro area and introduction of the euro.
- January 1, 2023: Dual circulation begins.
- January 15, 2023: Dual circulation ends (exchange of kuna cash for euros and withdrawal of Croatian kunas from circulation). The euro becomes the sole legal tender.
- January 15, 2023: End of dual pricing.

Bulgaria

- January 1, 2007: EU membership.
- July 10, 2020: ERM II (European Exchange Rate Mechanism II) membership and entry into close cooperation with the ECB.
- July 8, 2025: Council of the EU decision approving Bulgaria's accession to the euro area and fixing the conversion rate at 1.95583 lev per €.

- August 8, 2025: Dual pricing in Bulgarian lev (BGN) and euro (€) starts.
- Summer–Autumn 2025: Euro banknote and coin production and frontloading to the Bulgarian National Bank and commercial banks.
- November–December 2025: Sub-frontloading of euro cash to retailers and distribution of starter kits to the public.
- January 1, 2026: Accession into the euro area and introduction of the euro as the official currency.
- January 1, 2026: Dual circulation begins – the lev and the euro are both legal tender.
- January 31, 2026: Dual circulation ends – the euro becomes the sole legal tender; lev banknotes and coins can still be exchanged at the Bulgarian National Bank without time limits.
- December 31, 2026: End of dual pricing in BGN and €.