Has the Euro Shrunk the Band?
Relative PPP Convergence in a Currency Union

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Abstract

I study the effects of entry to the European Monetary Union (EMU) on relative PPP convergence using monthly disaggregated price indexes from 32 European countries from 1999 to 2016. I apply Heckscher's insight that transaction costs create bands of inaction in which price differences are not arbitraged away. I examine the entry of Cyprus, Malta, Slovakia, and Slovenia to the EMU, and estimate the bands of inaction before and after entry using a threshold autoregressive model. I find a positive effect of the EMU on relative PPP convergence: after entry the bands of inaction with EMU members fell by 17%. Cross sectional evidence further confirms the result and supports the theoretical prediction that bands of inaction are related to transaction costs.

Keywords: Currency Union, Relative PPP Convergence, Euro, Arbitrage, Threshold Autoregressive Model.

JEL Code: F15, C24.

*lmacedoni@econ.au.dk. I am grateful to Alan Taylor for suggestions and support. I also thank Charles Engel, Deborah Swenson, Ina Simonovska, three anonymous referees, and seminar participants at UC Davis International/Macro brown bag, and Western Social Science Association Conference 2017.

This article has been accepted for publication and undergone full peer review but has not been through the copyediting, typesetting, pagination and proofreading process, which may lead to differences between this version and the Version of Record. Please cite this article as doi:

10.1111/sjoe.12417

1 Introduction

Does a currency union improve price convergence among of its members? In theory, yes: a common currency reduces transaction costs and, hence, facilitates price convergence (Mundell, 1961). The twenty years of the EMU have provided a relevant experiment to test such an intuitive theory. However, the evidence is mixed. In fact, while deviations from PPP are smaller within the EMU (Cavallo et al., 2014), early studies often found that price convergence occurred before the birth of the currency union (Engel and Rogers, 2004; Rogers, 2007).[1] However, these early studies had to examine the first few years of the EMU, whose founding members began the process of economic integration after the Second World War. As the EMU expanded, it is now possible to examine the entry of countries lacking a prior history of economic integration with the other EMU members.

To account for the effect of the EMU on price convergence, this paper examines the entry of Slovenia, Cyprus, Malta, and Slovakia to the EMU between 2007 and 2009. I use a comprehensive data of monthly disaggregated price indexes from Eurostat, which span from the introduction of the euro in January 1999 to March 2016. I estimate the bands of inaction for the relative prices of 43 tradable baskets of goods between the four entrants and the other European countries before and after the entry in the EMU. Bands of inaction are defined by Obstfeld and Taylor (1997) as "*thresholds delineating a region of no central tendency among relative prices*". I find a positive effect of the EMU on relative PPP convergence: the bands of inaction for the country pairs that started sharing a common currency have shrunk.

The theoretical background for the estimation of bands of inaction traces back to Heckscher (1916), who noted that transaction costs create scope for deviations from the law of one price (LOP) and, hence, from purchasing power parity (PPP). I formalize the insight in a simple model of international arbitrage, in which arbitrageurs must pay variable and fixed costs to exploit international price differences for profit. The model predicts that the bands of inaction between two countries increase with variable and fixed transaction costs, and decline with the size of the two countries. By reducing transaction costs, a currency union shrinks the bands of inaction and, thus, it facilitates price convergence. Such an effect is reflected by smaller bands for price indexes.

I estimate the thresholds of the bands of inaction of the deviations from PPP using a threshold autoregressive (TAR) model (Obstfeld and Taylor, 1997). For each time period (before and after the entry to the EMU), I estimate good and country pair specific thresholds of the band of inaction. Across periods, the average band of inaction is ±12%: any differences between the relative price of a basket of goods and its long-run equilibrium value outside the 12% range would be arbitraged away. I find that entry of Cyprus, Malta, Slovenia, and Slovakia to the EMU is associated with

[1] Cavallo et al. (2014) compare online price dispersion within the EMU and outside the EMU while in the present study the object of analysis is the change in price dispersion after entry to the EMU. The more recent work by Dvir and Strasser (2018) finds that the EMU improves price convergence in the car market, but that most of the convergence occured prior to the EMU creation. The present paper examines a wider range of products and documents no prior price convergence for the new members of the EMU.

a reduction in the thresholds of the bands of inaction: the bands of inaction for the country pairs that started sharing a common currency fell by 17%. Furthermore, before entry, bands of inaction were not significantly lower with respect to future EMU members, which indicates that the integration between the four entrants and the EMU members was no greater than their integration with European countries outside of the EMU.

Countries that are members of the EMU are also members of the European Union (EU) and, for the most part, of the Schengen Area. Both the EU and the Schengen Area are European institutions that promote economic integration. In fact, the average band of inaction is ±11% within the EU and ±15% outside the EU. Similarly, country pairs in the Schengen Area have an average band of inaction of ±11% while outside of the Schengen Area, the average band is ±13%. To identify the effect of the EMU separately from these other European institutions, I restrict the sample of countries to members of the EU and the Schengen Area. Conditional on this restricted sample, the result are robust: the band shrunk by 25%.

The baseline TAR model assumes that the thresholds of the bands of inaction are constant within each period (before and after entry). However, it is reasonable to assume that bands of inaction vary over time, to reflect technology shocks or institutional changes that improve economic integration over time. To account for this possibility, I consider an alternative TAR model where the thresholds of the bands have a linear trend. Hence, I can study how the intercept, or the initial level of the threshold, and the slope of the trend change after entry to the EMU. First, I do not find any significant change in the intercept: the initial level of the threshold with respect to EMU members remains constant after entry, relative to non EMU members. This result confirms the lack of better integration between the four entrants and EMU members between 1999 and their entry dates, relative to the other European countries. Second, I find that the slope between entrants and EMU members, relative to other countries, is smaller after entry, which indicates that economic integration improved only after entry.

As a complement to the previous result, I integrate my dataset with the price indexes for selected goods provided by Juvenal and Taylor (2008) from 1981 to 1998. The sample of countries contains the founders of the EMU and two countries that opted out of the common currency: Denmark and the United Kingdom. Although the small sample of countries suggests caution in the interpretation of the results, the findings are quantitatively large. After the creation of the EMU in 1999, the bands of inaction of country pairs that shared the same currency fell by approximately 60% relative to the country pairs that kept different currencies. Contrary to the case of Cyprus, Malta, Slovakia, and Slovenia, examining the time varying thresholds suggests that some of the recorded integration was already present before the creation of the EMU. This means that economic integration among the EMU founders begun before entry to the EMU, which is in line with the results of Goldberg and Verboven (2005) and Rogers (2007).

As a robustness check, I conduct a cross-country examination of the thresholds of the bands of inaction for the period 1999 to 2016. This analysis confirms the negative euro effect on the bands of inaction: country pairs that share the same currency have bands of inaction that are 14% smaller than the average band. Furthermore, the cross-country analysis allows me to compare the magnitude of the EMU effect relative to other transaction costs. Controlling for the volatility of the nominal exchange rate, I find that the effect of the euro on the bands of inaction exceeds that of a fixed exchange rate regime, which confirms the results of Cavallo et al. (2014).

The cross-country analysis confirms the theoretical prediction that bands of inaction are related to transaction costs. Bands of inaction are wider for more distant country pairs and for islands and narrower for country pairs that share a common language. Moreover, the evidence indicates the presence of fixed costs of international arbitrage: the larger the countries are, the lower the transaction costs between them are. There is a negative relationship between bands of inaction and the dimension of the distribution network, proxied by measures of the size of retail, wholesale, and online markets. Finally, I estimate the bands of inaction for non-tradable goods: neither the euro nor bilateral distance affects the bands of inaction of non-tradable goods.

However, the model predicts that bands of inactions for relative price indexes are an indirect measure of the underlying transaction costs. First, the bands of inaction of price indexes are a fraction of the bands of individual goods, and the fraction is proportional to the expenditure shares on each good. Second, since arbitrage usually occurs through non-traditional distribution channels, such as parallel imports (Ganslandt and Maskus, 2004), there is a certain degree of substitutability between goods sold in traditional channels and the goods sold by arbitrageurs. Such substitutability dampens the effects of transaction costs on bands of inaction. For these two reasons, the estimated effect of the euro on the bands of inaction can be considered as a lower bound for the effect of the euro on transaction costs.

What types of transaction costs does the EMU reduce? First, the EMU eliminates the need to hedge against exchange rate risk, which is analogous to fixed exchange rate regimes. However, the results of the study imply that the effect of the EMU goes beyond that of a pegged system. Second, a common currency reduces currency exchange fees (Anderson and Wincoop, 2004). Third, a currency union reduces information frictions by facilitating the comparison of prices. In fact, experimental evidence from Mussweiler and Strack (2004) suggests that consumers find it easier to compare prices in the same currency.[2]

In the last few years, the European crisis (Lane, 2012) and the recent vote on the exit of the UK from the EU cast some shadows on the European experiment. This study addresses both issues. First, critics of the EMU claim that in the presence of country-specific shocks, the lack of exchange rate flexibility limits price adjustments. However, the results of this study partially contrast this claim for tradable goods: within the EMU, bands of inaction are narrower, and relative PPP convergence is faster when it is outside the band of inaction, in line with Bergin et al. (2017).

[2] Boivin et al. (2012) argue that information frictions can explain the deviations from the LOP between the US and Canada, despite minimal transaction frictions. A related mechanism is that of Cavallo et al. (2014), who argue that firms tend to charge similar prices in a currency union to avoid angering their customers (Rotemberg, 2005).

Moreover, countries that are in the EU exhibit bands of inaction that are lower than the average band. The result implies that arbitrage will become more expensive and, thus, prices will likely diverge between the UK and the members of the EU.

The remainder of the paper is organized as follows. Section 2 surveys the related literature. Section 3 presents a simple model of arbitrage that guides the empirical analysis. Section 4 describes the data and it outlines the empirical strategy to estimate the bands of inaction. Section 5 provides evidence for the negative relationship between the euro and the bands of inaction. Section 6 concludes.

2 Related Literature

Several studies have examined the effects of the EMU on price convergence, finding mixed results. Engel and Rogers (2004), Parsley and Wei (2008), and Fischer (2012) are among the studies that reject the hypothesis of price convergence within the EMU. By contrast, Goldberg and Verboven (2004), Lutz (2004), Allington et al. (2005), Dvir and Strasser (2018), Cavallo et al. (2014), and Cavallo et al. (2015) provide evidence for a positive effect of currency unions on price convergence.[3]

The dataset and the methodology used distinguish this study from the previous literature. The main benefit of the dataset is the period covered, which spans from 1999 to 2016, and thus includes the entry of Slovenia, Cyprus, Malta, and Slovakia in 2007, 2008, and 2009. With the exception of Cavallo et al. (2015), prior research on price convergence in the EMU could not include new entrants to the EMU — merely because they were not in the EMU yet. Hence, studies on the effects of entry to the EMU, such as Engel and Rogers (2004), Allington et al. (2005), Parsley and Wei (2008), and Fischer (2012) had to examine changes in price dispersion between the founders for a few years after their 1999 EMU entry. As the founders of the EMU had a long history of economic integration, it is no surprise that most authors found that price convergence occured years or decades prior to the EMU (Goldberg and Verboven, 2005; Rogers, 2007; Cuaresma et al., 2007; Faber and Stokman, 2009). In contrast, I verify the lack of improved economic integration prior to entry of the four countries mentioned, and examine the change in their bands of inaction after their entry.

Examining the variation in bands of inaction after entry to the EMU allows to avoid the common problem of endogeneity in the entry to a currency union (Alesina and Barro, 2000). In fact, evidence from the cross-country comparison of price convergence such as those of Lutz (2004) and Cavallo et al. (2014), cannot establish whether a currency union causes price convergence, or whether price convergence was achieved prior to entry to a currency union.

Another benefit of the dataset is its product coverage. Using disaggregated price indexes for European countries makes it possible to estimate bands of inaction for a wide set of products. It

[3] A related research question is whether a currency union promotes trade. In the context of the EMU, the currency union effect on trade flows is in the range of 4-50% (Micco et al., 2003; Baldwin, 2006; Flam and Nordström, 2006; Glick and Rose, 2016).

allows for general conclusions on the relationship between prices and currency unions relative to studies that focus on a single good or a small set of products. This is particularly important as the euro effects appear to be heterogeneous across goods. Examining a similar sample of countries of years, Fischer (2012) and Parsley and Wei (2008) find no euro effect on the prices of washing machines and the Big Mac and its ingredients. In contrast, Goldberg and Verboven (2004) and Cavallo et al. (2015) document a euro effect on the prices of cars and online retailers.

Deviations from LOP and PPP are not only a cross-country phenomenon, but they are also present within a country. Engel and Rogers (1996) were the first to estimate trade costs using price data, by comparing price variation across cities within US and Canada, and across the two countries. Deviations from LOP across cities within a country are documented in Japan (Crucini et al., 2010), US (Handbury and Weinstein, 2015), and Canada (Hickey and Jacks, 2011). In addition, several studies examined convergence to the LOP as a measure of improved market integration, from an historical perspective. For instance, Crucini and Smith (2016) find that the effect of distance on deviations from LOP within Sweden declines from 1732 to 1914, and Dobado and Marrero (2005) document convergence in corn prices across Mexican states from 1876 to 1910.

I use the TAR model (Tong and Lim, 1980) to estimate the bands of inaction of relative price indexes.[4] I borrow the TAR model from the literature that argues that non-linearities in the time series of real exchange rates explain deviations from LOP and PPP.[5] This study is closely related to Parsley and Wei (2008) who estimate a TAR model using data on the prices of a Big Mac and its ingredients from 1993 to 2006. The authors assume that the EMU only affects the speed of price reversion, and that the bands of inaction are good specific and are constant across country pairs. In this study, the variable of interest is the band of inaction, which is assumed to be good and country pair specific. I find that the EMU reduces the bands of inaction and, furthermore, increases the speed of convergence.

The bands of inaction provide an indirect measure of transaction costs and, thus, the results of this study suggest that the joining the EMU is associated with lower transaction costs. Despite deviations from the LOP and PPP imply the existence of transaction costs, the literature lacks a systematic assessment of the extent of international arbitrage and its costs. Evidence of international arbitrage is limited to particular industries and distribution channels. A popular case in the economic and public health literature is the smuggling of cigarettes. For instance, Galbraith and Kaiserman (1997) document that the presence of smuggled cigarettes limits the effect of taxes on cigarettes in Canada. Another type of arbitrage is the case of parallel imports, which are *"goods traded without the authorization of an original trademark or copyright owner"* (Maskus and Chen,

[4] Hansen (1997), Hansen (1999), and Caner and Hansen (2001) further develop the theory of TAR models.

[5] According to the TAR model, the relative price of the same goods in two locations is a random walk within a band of no-arbitrage while outside of the band the relative price is mean-reverting (Obstfeld and Taylor, 1997). The Exponential Smooth Transition Autoregressive model instead assumes that the relative price is always mean reverting but the speed of reversion is higher the higher the distance from the mean (O'Connell and Wei, 2002). Juvenal and Taylor (2008) use a Self-Exciting Threshold Autoregressive (SETAR) model, whose main difference with the TAR model is that the number of lags in the AR process is chosen according to the information criterion.

2002). A report from NERA (1999) illustrated that the share of parallel imports over industry imports ranges from less than 5% to 15%. A similar share is reported by Verboven (1996) in the automotive industry. Ganslandt and Maskus (2004) and Thompson (2008) document how prices are arbitraged through parallel imports in the case of pharmaceutical products and of digital cameras.

The rise of online retail had a twofold effect on transaction costs. First, the ease with which price information can be gathered online reduces the costs to arbitrage a good sold online and in traditional stores.[6] Second, online stores introduce virtual barriers that limit the websites accessible to customers: a consumer from the EU cannot make purchases on the iTunes market in the US. Anecdotal evidence suggests that arbitrage is still possible even in the presence of virtual barriers. For instance, iTunes US gift cards allow access to the US market, regardless of the geographic location of the consumer. As a result, outside the US, iTunes US gift cards are sold above face value on eBay (Ng, 2013). In this study, I find that the extent of online purchases has a negative effect on bands of inaction.

3 A Simple Model of International Arbitrage

This section builds a model of international arbitrage that shows how arbitrage costs, either variable or fixed, generate bands of inaction for the relative prices of goods. Since the dataset used in the empirical analysis contains information about the price of baskets of goods, I show how the bands of inaction for individual goods generate bands of inaction for the relative price of baskets of goods.

Consider a given basket of goods for two countries i and j. Goods are indexed by $b = 1, ..., B$. Let $\tilde{p}_{ibt}$ be the price of good b in country i and period t, and $\tilde{p}_{ib0}$ the same price at some reference period 0. Let the corresponding prices in country j be expressed in the same currency. Finally, the expenditure share of good b at some reference period 0 is w_{ib0}, with $\sum_{b=1}^{B} w_{ib0} = 1$. The empirical analysis uses HICP data for disaggregated baskets of goods, which are Layspeyres-type price indexes. Consistent with the HICP methodology, I define P_{it} as the weighted geometric average of prices for the basket of goods in country i and period t:

$$P_{it} = \prod_{b=1}^{B} \left(\frac{\tilde{p}_{ibt}}{\tilde{p}_{ib0}} \right)^{w_{ib0}} \tag{1}$$

Let $p_{ibt} = \ln \tilde{p}_{ibt}$ denote the log of prices. Taking the log of (1) yields $\ln P_{it} = \sum_{b=1}^{B} w_{ib0}(p_{ibt} - p_{ib0})$. Let RP_{ijt} denote the log difference between the price index in country i and in country j:

$$RP_{ijt} = \ln P_{it} - \ln P_{jt} = \sum_{b=1}^{B} (w_{ib0} p_{igt} - w_{jb0} p_{jbt}) - \sum_{b=1}^{B} (w_{ib0} p_{ib0} - w_{jb0} p_{jb0}) \tag{2}$$

RP_{ijt} is made of two summations terms. The first one represents the price differences across goods,

[6] See Alaveras et al. (2015) for the case of books. Moreover, there is a number of websites and books that claim to explain how to exploit arbitrage opportunities through Amazon.

with each price weighted by the corresponding expenditure share. The second term represents the cross-country price difference in the reference period.

Let RP^*_{ijt} denote the relative price indexes (2) evaluated in the case where the law of one price holds, namely $p_{ibt} = p_{jbt}$, $\forall b$.[7] This implies that all goods are sold in both countries or, alternatively, that LOP holds even for goods that are not available. RP^*_{ijt} equals:

$$RP^*_{ijt} = \sum_{b=1}^{B} p_{ibt}(w_{ib0} - w_{jb0}) - \sum_{b=1}^{B}(w_{ib0}p_{ib0} - w_{jb0}p_{jb0})$$

Since we are dealing with price indexes, RP^*_{ijt} does not equal zero when the law of one price holds. In fact, the first summation term represents the difference in the expenditure shares across countries while the second term captures the cross-country price difference in the reference period. Anticipating the procedure I follow in the empirical analysis, let us consider the deviations of RP_{ijt} from the long run equilibrium where the LOP, and, thus, PPP hold:

$$RP_{ijt} - RP^*_{ijt} = \sum_{b=1}^{B} w_{jb0}(p_{ibt} - p_{jbt}) \tag{3}$$

Deviations from the long run equilibrium of PPP are given by the weighted average of the log price difference between the two countries, or the deviations from LOP. The weights in the average are the expenditure shares in one country. To derive the bands of inaction for (3), let us first consider how arbitrage affects the relative price of a single good.

3.1 Arbitrage and Bands of Inaction

Consider an identical good b sold at different prices $\tilde{p}_i$ and $\tilde{p}_j$, expressed in the same currency, in two countries i and j. I drop product and time subscript for the sake of clarity.[8] Purchasing the good in i and selling it to j is subject to frictions. First, arbitrageurs pay a variable cost τ_{ij} that captures all variable costs involved with arbitrage (for instance, shipping and tariffs). Arbitrage also involves a set of activities whose cost is independent of the quantity shipped, such as fixed fees for storage, transport or currency exchanges. Moreover, legal barriers such as property rights, safety regulations and exclusive distribution channels create additional costs to arbitrage. I summarize these frictions as a fixed cost of arbitrage f_{ij}.

International arbitrage often occurs through non-traditional distribution channels, and in order to benefit from lower prices of arbitraged goods, consumers incur in switching costs. For instance, price arbitrage of books requires consumers to switch between traditional and online retail (Alav-

[7] Without loss of generality, I treat prices in country i as reference, or as the long run equilibrium of prices.

[8] If price-arbitrage requires time, uncertainty — about exchange rates or demand — would play a role. Arbitrage costs would be higher for country pairs that exhibit higher uncertainty. The presence of uncertainty suggests that the bands of inaction are wider for countries in a flexible exchange rate regime. The empirical analysis confirms such a prediction.

eras et al., 2015). Moreover, parallel imports are sold through distributors other than the licensed ones (Ganslandt and Maskus, 2004). I summarize those switching costs with the assumption that the good sold by traditional retailers and the good sold by arbitrageurs are imperfect substitutes. Let Q_j denote the demand for the good in country j and p_{ij} denote the price charged by the arbitrageur that purchases the good in i and sells it in j. Because of the switching costs faced by consumers, the arbitrageur is only able to capture a fraction of Q_j. In particular, the arbitrageur faces a demand q_{ij} equal to:

$$q_{ij} = Q_j \left(1 - \frac{p_{ij}}{\tilde{p}_j}\right)^{\gamma} \tag{4}$$

where the parameter γ captures the degree of substitution between the two goods. If $\gamma = 0$, the model boils down to the textbook model of arbitrage, where arbitrageur captures the entire demand by charging a price slightly below the market price.[9] Larger values of γ reduce the optimal price that an arbitrageur charges.

Because arbitrage entails the payment of variable and fixed costs, it is only profitable if the price difference between two countries is large enough. In particular, there exist two thresholds that delimit the band of inaction for the relative price of a good within which arbitrage does not occur. The band of inaction is defined as $\frac{\tilde{p}_j}{\tilde{p}_i} \in \left[\frac{1}{1+\underline{c}_{ij}}, 1 + \bar{c}_{ij}\right]$, where $\underline{c}_{ij}$ and $\bar{c}_{ij}$ denote the lower and upper thresholds. I leave the derivations to the online appendix.

The thresholds of the bands of inaction are given by:

$$\underline{c}_{ij} = \frac{\tau_{ji} + \frac{1+\gamma}{\gamma^{\frac{\gamma}{1+\gamma}}}\left(\frac{f_{ji}}{E_i}\right)^{\frac{1}{\gamma+1}} - 1}{1 - \frac{1+\gamma}{\gamma^{\frac{\gamma}{1+\gamma}}}\left(\frac{f_{ji}}{E_i}\right)^{\frac{1}{\gamma+1}}}; \qquad \bar{c}_{ij} = \frac{\tau_{ij} + \frac{1+\gamma}{\gamma^{\frac{\gamma}{1+\gamma}}}\left(\frac{f_{ij}}{E_j}\right)^{\frac{1}{\gamma+1}} - 1}{1 - \frac{1+\gamma}{\gamma^{\frac{\gamma}{1+\gamma}}}\left(\frac{f_{ij}}{E_j}\right)^{\frac{1}{\gamma+1}}}$$

where $E_j = Q_j\tilde{p}_j$ denote total expenditures on the good in country j. The width of the band of inaction increases in the variable cost of arbitrage τ_{ij}, which is the standard assumption in the literature (Obstfeld and Taylor, 1997; Anderson and van Wincoop, 2003). With $f_{ji} = 0$, the bands of inaction become identical to the textbook case. Despite the substitutability between original and arbitraged good, the bands of inaction would be defined by the two variable arbitrage costs. Fixed costs of arbitrage create a role for market size: the larger the size of both countries i and j, the lower the band. Moreover, the effect of variable and fixed costs on the band of inaction depends on the degree of substitution between the arbitraged good and the original good γ.[10] For the thresholds of the band to be symmetric, arbitrage costs need to be symmetric as well, and expenditures identical across the two countries.

[9] An exception to the textbook case is Maskus and Chen (2002), who assume that parallel importers and official retailers compete oligopolistically. The online appendix outlines the result under the assumptions of perfect substitution with and without fixed costs of arbitrage.

[10] The first order approximation of the threshold equals $\ln(1 + \bar{c}_{ij}) \approx \bar{c}_{ij} = \ln\tau_{ji} - \ln\left[1 - \frac{1+\gamma}{\gamma^{\frac{\gamma}{1+\gamma}}}\left(\frac{f_{ji}}{E_i}\right)^{\frac{1}{\gamma+1}}\right]$.

3.2 Relative Price Indexes and Bands of Inaction

Equation (3) shows that deviations from the long run equilibrium of PPP are given by the weighted average of the log price differences of each good in the basket. The log price differences of individual goods are affected by arbitrage as soon as they go beyond the commodity points. What are, then, the corresponding commodity points for relative price indexes?

To gather some intuition, consider the case in which all goods $b = 1, ..., B$ in the basket have the same threshold for the band of inaction $\bar{c}$, and the same weights w. In such a case, the band within which arbitrage does not occur is given by $\pm w\bar{c}$. To derive such a result, suppose for all but one good v LOP holds. The deviations from PPP are then equal to the log price difference for such good, weighted by its expenditure share in the basket. Namely, $RP_{ijt} - RP^*_{ijt} = w_{jv0}(p_{ivt} - p_{jvt})$. Arbitrage occurs when the log price difference $p_{ivt} - p_{jvt}$ is outside the thresholds $\pm\bar{c}$. When the log price difference is outside the threshold, $RP_{ijt} - RP^*_{ijt}$ is outside of the thresholds $\pm w\bar{c}$. Hence, the bands of inaction for relative price indexes are smaller in magnitude that the band of inaction for the relative price of each good. The larger the number of goods in the basket, the smaller the expenditure shares w, the larger the difference between commodity points for individual goods and for baskets of goods.

Let us now consider the case in which expenditure shares w_{jb0} and commodity points differ across goods $\bar{c}_b$. Let $c_{min} = \min_{b=1,...,B}\{w_{jb0}\bar{c}_b\}$ denote the lowest commodity point weighted by the corresponding expenditure share. Similarly, let $c_{max} = \max_{b=1,...,B}\{w_{jb0}\bar{c}_b\}$ denote the maximum. Deviations from PPP are not affected by arbitrage as long as the lay within the band $\pm c_{min}$. When the deviation from PPP lays between $[c_{min}, c_{max}]$ and between $[-c_{max}, -c_{min}]$, arbitrage may or may not occur, depending on whether the individual goods are beyond their respective thresholds. Finally, if $RP_{ijt} - RP^*_{ijt}$ is outside of the band $\pm c_{max}$, arbitrage will occur and relative price indexes will converge back towards the band.[11]

When two countries enter a currency union, the bands of inaction of their relative prices could shrink, because either variable costs τ_{ij} or fixed costs f_{ij} fall. If the band of inaction of relative price shrink, we would observe a reduction in the bands of inaction for relative price indexes as well. However, given a change in the bands of inaction for relative prices, the corresponding reduction in relative price indexes would be smaller, as it would be weighted by the expenditure shares of each good.

There are several channels through which a common currency could facilitate international arbitrage. First, as in a fixed exchange rate regime, hedging currency risk is no longer necessary in a currency union. Second, a common currency eliminates fees and other costs related to currency exchange and settlement delays (Allington et al., 2005). Finally, there could be a psychological component related to a common currency: the results of Mussweiler and Strack (2004) suggest

[11] In the presence of heterogeneity in commodity points and expenditure shares across the goods in a basket, the estimation procedure I describe in the following section returns an average of the thresholds of the bands of inaction of each individual goods. Details are in the online appendix.

that a common currency increases consumers' ability to compare prices.[12]

4 Empirical Strategy

Guided by the model, this section tests whether a currency union reduces the bands of inaction for the relative price indexes of disaggregated baskets of goods. I begin by describing the sources of data. Then, I present the empirical methodology to estimate the bands of inaction.

4.1 Data

I use data on the Harmonized Index of Consumer Prices (HICP) of 80 disaggregated baskets of goods for 32 European countries from January 1999 to March 2016. A basket is a four-digit good of the Classification of Individual Consumption According to Purpose (COICOP). I divide the baskets into tradable and non-tradable following the division proposed by Sturm et al. (2009).[13] The main results of the study are derived from the sample of 43 tradable four-digit COICOP baskets. The appendix reports the list of countries and the dates of their entry to the EMU, as well as the list of goods considered in the analysis.[14]

A potential concern is that the entry in the EMU corresponds to changes in the methodology used to collect price data. However, all the countries in the sample started collecting price data according to HICP methodology before January 1999 or in the early 2000's. Moreover, Eurostat conducts regular compliance monitoring activities to verify that countries in the EU adhere to the guidelines, recommendations, and methodology of HICP. All the national practices of data collections of EU countries inspected between 2006 and 2016 are deemed comparable and of adequate precision and representativity. The online appendix summarizes the results of the monitoring activities.

The main issue of using HICP to estimate the thresholds of the bands of inaction is that price indexes do not carry cross-country information. Thus, this study mainly focuses on relative PPP convergence and the model showed the relationship between relative price convergence and convergence of the deviations from PPP.

There are few additional limitations. First, there is a possible downward aggregation bias of price dispersion (Allington et al., 2005): the dispersion of prices of baskets of goods is smaller than the dispersion of prices of individual goods.[15] To understand it, consider a basket of two goods,

[12] Because of this psychological reason, Cavallo et al. (2015) argue that firms reduce price differences of their products across countries to avoid angering their customers.

[13] Such a division may be affected by trade costs: non-tradable goods are such because their trade costs are prohibitive. The endogeneity of the division of product categories into tradable and non-tradable is less of a concern here, since Sturm et al. (2009) categorize goods according to product description. Most non-tradable are services of various types such as repair, transport, and insurance.

[14] The frequency of price observations may influence the estimation of the bands of inaction. Using high frequency prices may lead to wider bands of inaction than low frequency prices, due to the time required for arbitrage.

[15] Using intranational price data from Denmark, Michael et al. (2018) document that the estimated half-lives of relative prices are highly biased by aggregation procedures.

with the same price index in France and Germany. Suppose that one of the goods is cheaper in France and the other good is cheaper in Germany. Although consumers arbitrage away the price difference of the two goods, the price indexes remain unchanged. As a result, using price indexes underestimates the bands of inaction and, thus, the costs of international arbitrage. Results from simulations suggest that the aggregation bias is mild and depends on the number of products inside the basket of goods considered. The bias is strong for a small number of products, but it becomes weaker for larger number of products.

A second limitation follows from the use of price indexes (3). Assuming, for the sake of exposition, that all goods within a basket have the same band of inaction, the estimated band for the basket of good will be a fraction of the bands of individual goods. Such a fraction is proportional to the expenditure share on each good, and it is, thus, decreasing in the number of goods. Therefore, the larger the number of goods within a basket, the smaller the band of inaction of the basket.[16]

As national agencies decide the number of goods within each basket, and such a number is relatively small, the negative relation between number of goods and bands of inaction might be mild. To see why, let us examine further how the HICP methodology computes price indexes. Within each four-digit COICOP code there is a number of products that varies by country. Examples of products are: crackers, bread sticks, durum wheat pasta, egg pasta etc... Within each product, national agencies choose the most representative items. On average, there are 700 products across countries: nine products for each of the four-digit COICOP code I consider in the baseline case. Although it is possible to find information on the number of products for a few countries, the number of items is unlikely to be public, let alone the specific details of each item. An exception is Italy where there were 1.484 items, grouped in 400 products in the 2016 HICP. This is equivalent to 19 items per four-digit COICOP code.[17]

New products and items are added and dropped every year to reflect changes in the consumption basket. These changes are, however, limited to a small number of items within the disaggregated baskets of goods. For Italy, there is an increase in the number of items over the years. For the earliest year available, 2004, there were 1.031 items, which correspond to 13 items per four-digit COICOP code on average. Over 12 years, the average number of items has increased by 6 units: one additional item per four-digit COICOP code every two years.

Despite the likely mild effect, both the aggregation bias and the negative relationship between estimated bands and number of goods are likely to downwardly bias the estimate of the bands and

[16]The online appendix provides the results of simulations to further discuss the two limitations. A third possible issue would arise if the relative price of goods within a basket were subject to common shocks. In such a case, the algorithm would estimate thresholds that are closer to the individual thresholds.

[17]The information can be retrieved on https://www.istat.it/it/archivio/179355 (in Italian). There might be variation in the frequency of items across products, but that information is not available. To the extent that the number of items is correlated to the number of products, food and non-alcoholic beverages have the largest number of products in Italy (120). This might explain why the estimated thresholds for the bands of inaction for food products are surprisingly small, given their perishable nature, when compared to other industries.

of arbitrage costs. A potential concern may arise if the aggregation bias were stronger in EMU countries relative to countries that do not share the same currency. Similarly, the results on the effects of the euro on the bands may not reflect a reduction in the costs of arbitrage but simply an increase in the number of products.[18] A possible test for the two concerns is the use of more aggregate price indexes. In fact, given the two biases, using more aggregate price indexes would magnify the effects of the euro on the bands of inaction. However, studying the effects of the euro on the bands of inaction for more aggregate price indexes (three-digit COICOP) yields results that are quantitatively similar to the baseline case. Such an exercise suggests that the two biases may only affect the level of the bands but not the effect of the euro on the bands.

4.2 Estimating the Band of Inaction

Let P^g_{it} be the logarithm of the harmonized index of consumer prices of the basket of goods g in country i at month t. Let z^g_{ijt} be the basket specific (log) price deviation from PPP between countries i and j, defined as

$$z^g_{ijt} = P^g_{it} - P^g_{jt} - s_{ijt} \tag{5}$$

where s_{ijt} is the logarithm of the nominal exchange rate, defined as the price of one unit of j's currency in terms of i's currency. The source of data for the monthly exchange rate is Eurostat. For each country pair ij and each basket g, I regress the log price deviation on a constant α^g_{ij}:

$$z^g_{ijt} = \alpha^g_{ij} + x^g_{ijt} \tag{6}$$

and record the error term x^g_{ijt}, which is the demeaned log deviation from PPP. In the theory section, such a concept is described by equation (3). Demeaning eliminates the prices evaluated at the reference period, as shown in (2).[19]

Let c^g_{ijt} be the threshold of the band of inaction for good g between country i and j at time t. The relatively limited number of observations does not allow me to estimate a threshold for each month t, nor for each year. Hence, I consider the following linear trend for the threshold $c^g_{ijt} = a^g_{ij} + b^g_{ij}(t-1)$. In my baseline specification, I set $b_{ij} = 0$, namely $c^g_{ijt} = c^g_{ij}$, which is the specification of (Obstfeld and Taylor, 1997). I estimate the thresholds under the two specification by maximum likelihood using the following TAR model:

[18] The number of different products available for consumption declines in trade costs and increases in market size (Hummels and Klenow, 2005). Thus, if the euro reduces arbitrage costs and trade costs, we may expect a reduction in the estimated bands through both a reduction in the bands of inaction for individual product c and a reduction in the expenditure shares on each good w.

[19] The procedure outlined by (Obstfeld and Taylor, 1997) additionally requires to detrend the log deviations from PPP. In such a way, the analysis abstracts from long run trends in relative price indexes and focuses on the short run adjustments toward the long run equilibrium. Moreover, the possible presence of trends in relative price indexes due to differentials in per capita income growth and monetary policy, may overestimate the effects of the euro on the bands of inaction. I have verified that the results are robust to this alternative specification. Details are in the online appendix.

$$\Delta x_{ijt}^{g} = \begin{cases} \lambda_{ij}^{g}(x_{ij\,t-1}^{g} - c_{ijt}^{g}) + e_{ijt}^{g,out} & \text{if } x_{ij\,t-1}^{g} > c_{ijt}^{g} \\ e_{ijt}^{g,in} & \text{if } c_{ijt}^{g} \geq x_{ij\,t-1}^{g} \geq -c_{ijt}^{g} \\ \lambda_{ij}^{g}(x_{ij\,t-1}^{g} + c_{ijt}^{g}) + e_{ijt}^{g,out} & \text{if } -c_{ijt}^{g} > x_{ij\,t-1}^{g} \end{cases} \tag{7}$$

where $e_{ijt}^{g,out} \sim N(0, \sigma_{ij}^{g,out2})$, $e_{ijt}^{g,in} \sim N(0, \sigma_{ij}^{g,in2})$. When the relative price is within the band of inaction ($|x_{ij\,t-1}^{g}| < c_{ijt}^{g}$), the relative price is a random walk. There are no forces of arbitrage inside the band of inaction, and thus, the best prediction of tomorrow's relative price is today's relative price. Outside the band, when $|x_{ij\,t-1}^{g}| > c_{ijt}^{g}$, the relative price is an AR(1) process, with coefficient λ^{out}. The larger the absolute value of λ^{out} is, the faster prices converge to the thresholds of the band of inaction $(+c_{ijt}^{g}, -c_{ijt}^{g})$. I estimate the band of inaction by a gridsearch algorithm, and compute the standard error of the estimated bands of inaction by bootstrap method, as suggested by Caner and Hansen (2001). The details are in the appendix.[20]

The empirical model assumes that the upper and lower thresholds of the bands of inaction are symmetric $(+c_{ijt}^{g}, -c_{ijt}^{g})$, which implies, according to the model, that fixed and variable costs of arbitrage, and the expenditures of each destination are symmetric too. In the online appendix, I explore this possibility by estimating asymmetric thresholds of the bands of inaction $(+\bar{c}_{ijt}^{g}, -\underline{c}_{ijt}^{g})$, finding similar effects of the euro. Furthermore, consistent with Waugh (2010), I find suggestive evidence that larger economies incur in smaller costs when arbitraging their cheaper products.

Another assumption of the empirical model is that the autoregressive coefficients λ_{ij}^{g} is the same whether the relative prices are above or below the bands. However, we might expect that country's characteristics, such as size, would influence the speed of reversion to the band. In the online appendix, I allow for different autoregressive coefficients depending on whether relative prices are above or below the band. Due to data limitations, the number of observations falls by 25%, and the coefficient are less precisely estimated. In fact, only in 8 % of the goods and country pairs, I fail to reject the hypothesis that the two autoregressive coefficients are different from one another. Despite these limitations, there is suggestive evidence that price convergence is faster when prices are higher in larger economies: larger demand facilitates price convergence.

5 Results

This section shows the main results of the study: when a country enters a currency union, the bands of inaction of relative price indexes between the country and the other members of the currency union shrink, relative to the countries with different currencies. First, I study the time variation in the bands of inaction before and after entry to the EMU. This analysis focuses on the

[20]The average bootstrapped standard error of the thresholds across tradable goods and country pairs for the full sample of years is 0.046, with a standard deviation of 0.03. Recall that the average threshold is 0.12.

entry of Cyprus, Malta, Slovakia, and Slovenia in the EMU. Furthermore, I consider the effects of the euro on the bands of inaction upon the creation of the EMU. Second, I study the cross-country variation in the bands of inaction using the entire sample available, which allows me to compare the relationship between EMU and bands of inaction to other proxies for transaction costs justified by the theoretical model.

5.1 Descriptive Statistics

Before examining the change in bands of inaction after the entry of countries to the EMU, it is useful to examine the descriptive results on the estimated bands of inaction. The results in this section rely on the entire sample of data provided by Eurostat, from 1999 to 2016.[21] The average threshold across tradable baskets of goods and country pairs is ±11.8%. In other words, any (absolute) differences between the relative price of a good and its long-run equilibrium value above 11.8% would be arbitraged away. Thresholds vary across baskets, suggesting that at least a component of costs to international arbitrage is good specific. Materials for the repair of the dwelling and non-durable household goods have the lowest average thresholds (8%) while information-processing equipment and photographic equipment have the highest (25%).

Previous studies estimate thresholds that are of similar magnitude. Obstfeld and Taylor (1997), using the US as a reference country and a more aggregated product classification, estimate thresholds in the range of ±7–10%. On the other hand, the commodity points estimated by Juvenal and Taylor (2008) are larger than those found here: using the US as the reference, the average commodity point the authors find with the SETAR model is ±17%. Finally, my estimates are larger than those of Parsley and Wei (2008), whose commodity points are common across countries and range between ±0.8 and ±6%.

The average threshold of the band of inaction is ±10.4% for country pairs that are in the EMU and ±12.5% for country pairs that do not share the same currency. Moreover, examining the baskets of goods, 37 out of 43 baskets exhibit a lower average band of inaction for country pairs that are in the EMU. For durables for recreation, meat, and spirits, the band of inaction between countries in the EMU is almost half that between countries not sharing the same currency.[22]

The descriptive analysis also suggests that other agreements and institutions — the EU and the Schengen Area — that promote economic integration in Europe, are related to smaller bands of inaction.[23] The average threshold c_{ij}^{g} for tradable goods is ±10.8% inside and ±15% outside the EU; ±10.9% inside and ±13.2% outside the Schengen Area, and ±10% inside and ±13.4% outside

[21] I also provide descriptive results in the following section, where I examine the change in thresholds for a few selected countries. However, using the entire sample of data allows for a larger number of observations for each good and country pair. Hence, the entire sample is more representative to examine average statistics, and each threshold is more precisely estimated.

[22] The online appendix reports the average bands of inaction at the product level, inside and outside the EMU.

[23] The result is important to keep in mind in the following sections, as all EMU countries are part of the EU, and Ireland and Cyprus are the only countries in the EMU that do not belong to the Schengen Area. Table 6 in the appendix provides the entry dates in these institutions for the countries in the sample.

both the EU and the Schengen Area.

Let us now consider the empirical specification where the thresholds have a linear trend. The average estimated coefficient on the trend is close to zero (10^{-6}), and, thus it is not a surprise that the average initial threshold for tradable goods is almost identical to the average constant threshold ($\pm$ 11.8%). To get a meaningful interpretation of the average slope, I multiply the slope by the number of months in the sample (219) and get and increase in the average threshold by 0.02% from $\pm$ 11.8% to $\pm$ 11.82%. Across the sample, the average constant threshold is highly correlated with the initial value of the time varying threshold (76.5%). In contrast, there is a negative correlation between initial threshold and slope (-37.5%), and a positive correlation between constant threshold and slope (31%). Thus, higher initial thresholds across goods and country pairs are associated with smaller slope, while higher average thresholds with larger slope.

Examining the average slope by basket of goods allows us to identify patterns in the evolution of good specific transaction costs between 1999 and 2016. Vegetables, garments, and furniture have the smallest average slope, with a reduction of the initial threshold of 7%, 3.5%, and 3%. In contrast, the largest slope are recorded for dairy products, photographic equipment, and information processing equipment, with an average increase of the initial threshold of 2.2%, 3%, and 11%. Only in 16 out of 43 products the average slope is lower between country pairs in the EMU than it is between country pairs not sharing the same currency, and especially so for information processing equipment, and liquid and solid fuels.

The estimated thresholds of the bands of inaction do not satisfy the triangle inequality in just 17% of the good and country-pair observations. The baskets of goods with the largest shares of violation are plants (19%) and photographic equipment (19%). In contrast, the smallest shares are recorded in recording media (6%) and liquid fuels (6%). Violations of triangle inequality are spread across countries in a somewhat equal manner. The country pairs Iceland-Czech Republic and Iceland-Poland have the largest shares of violation (46% and 43%), while there are several country pairs with no violations.

5.2 Evidence from the Time Series

To evaluate the effects of the EMU on the bands of inaction, the ideal test would compare, for each country that enters the EMU, the bands of inaction before and after the entry to the EMU. The control group for such test would be the country pairs whose currency did not change. However, for the estimation of the bands of inaction to be accurate enough, only countries that entered the EMU in the middle of the period 1999-2016 have enough observations prior and after the entry. Thus, my main empirical result relies on the entry to the EMU of the following four countries: Slovenia (in 2007), Cyprus and Malta (in 2008), and Slovakia (in 2009). Furthermore, as a robustness check, I combine my dataset with that of Juvenal and Taylor (2008), which spans from 1981 to 1998. For a selected number of countries and products, I can test the time series effect of the EMU on its founding members

The comparison of bands of inaction before and after the entry to the EMU allows me to tackle the issue of endogeneity in the choice of entering the EMU. If countries that decide to enter the EMU are already more integrated than those that decide to opt out or are preparing to enter the EMU, the currency union effect documented at a point in time might simply reflect the already existing lower transaction costs. Although anecdotal evidence by Rose (2001) suggests that the choice to join a currency union is not driven by a trade cost motive, the concerns are reasonable. In fact, Engel and Rogers (2004), Goldberg and Verboven (2005), and Dvir and Strasser (2018) document that price convergence in the EMU occurred long before the EMU was established.[24]

Furthermore, I am able to control for other sources of economic integration that are contemporaneous to the entry to the EMU. First, increased globalization, which might reduce transaction costs or increase the correlation of demand and supply shocks, would shrink the bands of inaction. Such a reduction is likely to be good specific, and thus affects the relative price of a good regardless of the currencies of a given country pair. In addition to that, other European institutions promote economic integration: both the EU and the Schengen area are correlated with smaller bands of inaction. By restricting the sample to country pairs within those institutions, I am able to isolate the EMU effect.

5.2.1 The Effects of Entry in the EMU

For each of the four countries that entered the EMU (Cyprus, Malta, Slovakia, and Slovenia), I divide the sample in two periods before and after their entry to the EMU. Let s denote the period, $s = 0$ if prior to the euro and $s = 1$ after the entry to the EMU. For each period and entrant, I estimate the two TAR models outlined in section 4.2 over the set of tradable baskets of goods. The procedure generates estimated constant thresholds of the bands of inaction c^g_{ijs} for the basket g between countries i and j in period s, where i = Cyprus, Malta, Slovakia, and Slovenia, and j denotes all other countries in the sample. Furthermore, in the model that allows for a time trend in the thresholds, I obtain the initial level of the threshold a^g_{ijs} and the slope of such trend b^g_{ijs}.[25]

The average band before entry to the EMU is $\pm 6.3\%$, and it ranges from $\pm 5.4\%$ for Malta to $\pm 7.5\%$ for Cyprus. After entry, the average band slightly declines to $\pm 6\%$: Cyprus and Slovakia experienced a reduction in the average band in the second period while the opposite occurred for Malta and Slovenia. The average band between the four entrants and the EMU members falls from $\pm 5.9\%$ to $\pm 5.2\%$ while the average band relative to the non EMU members remains constant at $\pm 7\%$.

[24]Rather than endogeneity, this could reflect the forward-looking behavior of agents, who anticipate the greater integration that occurs in a currency union (Bergin and Lin, 2012).

[25]For each entrant, I restrict the sample to have a symmetric time window around entry. For Slovenia, the pre-entry sample spans from 01:1999 to 12:2006, the post-entry sample spans from 01:2007 to 01:2013. For Malta and Cyprus, from 10:1999 to 12:2007 (pre-entry), and from 01:2008 to 03:2016 (post-entry). For Slovakia, from 10:2001 to 12:2008 (pre-entry), and from 01:2009 to 03:2016 (post-entry). Relative to cross-sectional analysis of the following section, the thresholds are less precisely estimated, since I use about half the observations.

For each country i, and sample period $s = \{0, 1\}$, I compute the log difference in the variable y^g_{ijs}, $\Delta y^g_{ij} = \ln y^g_{ij1} - \ln y^g_{ij0}$, for each country j and good g. $y^g_{ijs} = \{c^g_{ijs}, a^g_{ijs}, b^g_{ijs}\}$ takes on the value of the constant threshold of the band, the initial threshold, or the slope of the time trend in the time varying threshold case.[26] I then estimate the following regression:

$$\Delta y^g_{ij} = f_g + f_i + \beta \text{Euro}_{ij} + \epsilon^g_{ij} \tag{8}$$

where $\text{Euro}_{ij} = 1$ if the country pair ij shares the same currency after the entry to the EMU of i. f_g is a good fixed effect that captures any good specific change in the bands of inaction due to globalization or technology, f_i is a country fixed effect that captures the different entry of each country, and ϵ^g_{ij} is the error term. Table 1 shows the results using the log change in the constant threshold as dependent variable while table 2 uses the change in the initial threshold and in the slope of the time varying bands of inaction.

Table 1: Average Band and Entry to the EMU

	(1)	(2)	(3)
Euro	-0.173**	-0.172**	-0.250**
	(0.071)	(0.071)	(0.095)
Product FE	Yes	Yes	Yes
Entrant FE	No	Yes	Yes
Pairs in EU and Schengen	No	No	Yes
R^2	0.041	0.126	0.180
# Observations	3926	3926	2090

Results from OLS of equation (8). Robust std. error in parenthesis. Cluster: country pair. ***: significant at 99%, ** at 95%, * at 90%. Sample: tradable four-digit COICOP goods. Dependent variable: Δc^g_{ij}.

The coefficient on the euro dummy is negative and statistically significant at the 95% level: after entry to the EMU country pairs that started sharing the same currency have smaller bands of inaction relative to the country pairs that maintained two different currencies. The coefficient of -0.17 is interpreted as follows: entering the EMU reduces the bands of inaction by 17%. Since entering the EMU overlaps, for the i countries considered, with entry to the EU and Schengen Area, I condition the regression on the country pairs being both in the EU and Schengen Area.[27] The result are robust to this specification and the effect of the euro is slightly larger (25%).

The reduction in the bands of inaction suggests that the transaction costs associated with international arbitrage are smaller for members in the EMU. However, the reduction in the bands of 17% cannot be interpreted as a reduction in the transaction costs of 17%. In fact, the model showed that the estimated reduction in the bands of inaction provides a lower bound for the reduction in transaction costs: the bands of inaction for baskets of goods are only a fraction of the

[26] Due to the grid search algorithm, the slope can assume value of zero, which I drop. Furthermore, as the slope b^g_{ijs} can take on negative values, I use the simple difference $\Delta b^g_{ij} = 100 * (b^g_{ij1} - b^g_{ij0})$.

[27] This last regression drops Cyprus by construction.

bands for the individual goods in the basket. Moreover, since arbitraged goods are sold through parallel imports or unconventional distribution channels, they can be considered substitute for the goods sold through traditional channels. The substitutability of arbitraged goods farther dampens the effect of changes in trade costs on the bands of inaction.

Table 2: Time Varying Band and Entry to the EMU

	Initial Threshold			Slope of Time Varying Bands		
	(1)	(2)	(3)	(1)	(2)	(3)
Euro	0.113	0.113	-0.089	-0.023**	-0.023**	-0.013
	(0.074)	(0.074)	(0.077)	(0.010)	(0.010)	(0.015)
Product FE	Yes	Yes	Yes	Yes	Yes	Yes
Entrant FE	No	Yes	Yes	No	Yes	Yes
Pairs in EU and Schengen	No	No	Yes	No	No	Yes
R^2	0.062	0.140	0.173	0.009	0.071	0.080
# Observations	3707	3707	1995	3707	3707	1995

Results from OLS of equation (8). Robust std. error in parenthesis. Cluster: country pair. ***: significant at 99%, ** at 95%, * at 90%. Sample: tradable four-digit COICOP goods. In the first three columns, the dependent variable is the log change in the initial threshold of the band of inaction Δa^g_{ij}. In the last three columns, the dependent variable is the change of the coefficient of the time trend in the time-varying thresholds Δb^g_{ij}.

There is no clear pattern between entry to the EMU and the initial threshold of the bands of inaction with a linear trend. The coefficient on the Euro dummy is in fact statistically insignificant, and the point estimate takes on positive and negative values. The result is reassuring, as it indicates that, upon entry, the pairs that started sharing the euro did not start from a smaller threshold relative to the beginning of the previous period, relative to non EMU countries. Furthermore, the relationship between entry and slope is negative and statistically significant, with the exception of the sample that includes only pairs in the EU and Schengen Area. The result suggests that after entry, the currency union has improved relative PPP convergence over time.

There is considerable heterogeneity in the results at the basket level. I repeat the baseline regression (8) for each COICOP four-digit basket, and report the coefficients in the online appendix. For a large portion of products, I cannot find statistical significance, due to the relative small number of observations. Among those products for which results are statistical significant, the thresholds for durables for recreation and meat exhibited the largest reduction. The thresholds for oils and fats, soft drinks, materials for the repair of the dwelling, medical products, and electrical appliances fell by 50%. Furthermore, the slope of the time varying thresholds also fell for beer, non-durable household goods, and sport equipment.

Integration Prior to Entry. To verify the presence of economic integration prior to the entry to the EMU, I consider the cross-section relationship between bands of inaction and EMU, before and after entry. Guided by the theoretical predictions of the model, I use the following regression model in which the dependent variable is $y^g_{ijs} = \{c^g_{ijs}, a^g_{ijs}, b^g_{ijs}\}$:

$$y^g_{ijs} = \beta_1 \text{Euro}_{ij} + \beta_2 \tau_{ij} + \beta_3 \text{FER}_{ij} + \beta_4 \text{Expenditures}_{ij} + e_g + \epsilon^g_{ij} \tag{9}$$

The independent variables are

- $\text{Euro}_{ij} = 1$ if the pair ij was in the EMU at any time in my sample.
- τ_{ij} is a vector of the following proxies for trade barriers, from CEPII (Head et al., 2010):
  - $\text{Log(Distance)}_{ij}$ = log of bilateral distance between i and j;
  - $\text{Common Border}_{ij} = 1$ if i and j share a border;
  - $\text{Common Language}_{ij} = 1$ if i and j share an official common language or if a language is spoken by at least 4% of the population of i and j;
  - $\text{Island}_{ij} = 1$ if either i or j is an island;
  - $\text{Landlocked}_{ij} = 1$ if either i or j is landlocked;
- $\text{FER}_{ij} = 1$ if country pair in a fixed exchange rate regime or in the EMU.[28]
- Expenditures_{ij} = sum of the log of PPP adjusted real GDP in i and j, from the World Development Indicators.
- e_g = good fixed effect.

The geographical variables that I use are standard proxies for trade costs in the "gravity" trade literature (Head and Mayer, 2013). In the context of price dispersion, Lutz (2004) use a similar set of variables to study whether trade costs affect price dispersion. Finally, the existence of fixed cost of arbitrage motivates the inclusion of the variable for the size of country pairs, proxied by Expenditures_{ij}.[29]

The results are in table 3. The thresholds of the average bands of inaction do not exhibit any relationship with the euro prior to entry. However, after the entry there is a negative relationship between bands and euro. Relative to the average band of ±6%, bands are $.011/.06 = 18.3\%$ smaller. Let us now consider the time varying bands of inaction. The initial threshold between the four entrants and their future EMU partners is $.011/.081 = 13.5\%$ smaller than the average band. However, the slope is larger between (future) EMU members than it is between the other countries. After entry, the initial threshold relative to the average threshold barely changes for EMU members ($.007/.061 = 11.5\%$), but the slope significantly smaller for EMU members than it is for countries that do not share the same currency.

[28]Following Reinhart and Rogoff (2004), I divide country pairs into fixed and flexible regimes according to the median deviation of the bilateral exchange rate from its average in the sample. If the median deviation is below 5%, the dummy takes a value of one. The threshold of 5% is the largest crawling band allowed in Reinhart and Rogoff (2004).

[29]Although the model show that the effects of Expenditures_{ij} are non-linear (footnote 10), here I consider only the linear component.

Table 3: The Euro Shrunk the Band

	Constant Bands		Initial Threshold		Slope of Time Varying Bands	
	(Before)	(After)	(Before)	(After)	(Before)	(After)
Euro	-0.004	-0.011**	-0.019**	-0.007*	0.016*	-0.012***
	(0.005)	(0.005)	(0.008)	(0.004)	(0.009)	(0.004)
R^2	0.120	0.233	0.191	0.182	0.051	0.075
# Observations	3926	3926	3707	3707	3707	3707

Results from OLS of equation (9). Robust std. error in parenthesis. Cluster: country pair. ***: significant at 99%, ** at 95%, * at 90%. Sample: tradable four-digit COICOP goods. In the first two columns, the dependent variable is threshold of the band of inaction c_{ij}^g, before and after entry to the EMU. In the third and fourth columns, the dependent variable is initial threshold of the band of inaction a_{ij}^g, before and after entry to the EMU. In the last two columns, the dependent variable is the coefficient of the time trend in the time-varying thresholds $b_{ij}^g * 100$, before and after entry to the EMU.

We can interpret the results in the following way. Prior to EMU entry, the four countries examined were not relatively more integrated with their future EMU members than they were with the other countries. In fact, their average band for the pre-entry period is not statistically related to their future EMU membership. Furthermore, despite starting from a somewhat smaller threshold, the bands of inaction slightly widened with the future EMU members. Upon entry, the initial thresholds with the EMU members were at the same relative level with respect to the non EMU members. However, the time trend on the bands was significantly smaller within the EMU than outside. Hence, the average bands for the post-entry period are smaller within the EMU than outside.

5.2.2 The Effects of EMU Creation

To study the effects of the creation of the EMU, I use the data provided by Juvenal and Taylor (2008), which span from 1981:01 to 1998:12. The dataset covers the following nine countries: Belgium, Denmark, Germany, France, Italy, Netherlands, Portugal, Spain, United Kingdom, which are all members of the EU and Schengen Area. It provides price indexes for fourteen baskets of tradable goods: Alchohol, Books, Bread, Clothing, Communication, Dairy, Domestic Appliances, Footwear, Fruit, Fuels and Energy, Furniture, Meat, Sound and Photographic Equipment, Tobacco, and Vehicles. I match each good with the corresponding HICP category: some goods corresponds to four-digit COICOP goods, others to three-digit COICOP goods.[30]

I estimate $y_{ijs}^g = \{c_{ijs}^g, a_{ijs}^g, b_{ijs}^g\}$ for each country pair ij and basket of goods g. Let $s = 0$ denote the pre-EMU period, from 1981:10 to 1998:12, and $s = 1$ denote the post-EMU period, from 1999:01 to 2016:03. I estimate x_{ij0}^g using the data from Juvenal and Taylor (2008) and use

[30]In particular, I matched the Juvenal and Taylor (2008) basket *Alcohol* to *Alcoholic beverages* (COICOP 021), *Books* to *Books* (0951), *Bread* to *Bread and cereals* (0111), *Clothing* to *Clothing* (031), *Communication* to *Telephone and telefax equipment* (082), *Dairy* to *Milk, cheese and eggs* (0114), *Domestic Appliances* to *Household appliances* (053), *Footwear* to *Footwear including repair* (032), *Fruit* to *Fruit* (0116), *Fuels and Energy* to *Electricity, gas and other fuels* (045), *Furniture* to *Furniture and furnishings, carpets and other floor coverings* (051), *Meat* to *Meat* (0112), *Sound and Photographic Equipment* to *Audio-visual, photographic and information processing equipment* (091), *Tobacco* to *Tobacco* (022), and *Vehicles* to *Purchase of vehicles* (071).

Eurostat data for x^g_{ij0} both for COICOP three- and four-digit goods. Keeping in mind that the change in the methodology to compute price indexes might be biasing a simple comparison across averages, the average band of inaction prior to entry to the EMU was slightly smaller for the future EMU founders (±8.1%) relative to the countries that opted out (±9%). However, after entry the average bands of inaction fell to ± 5.8% for EMU pairs, and rose to ± 11.4% for non EMU countries. Details are in the online appendix.

I estimate (8), without the country fixed effect, as in this sample all countries enter the EMU at the same time. There are a few concerns in the interpretation of the results. First, there is a change in the methodology used to collect price data. During the nineties, all countries in the EU updated their price collection methodologies to the HICP guidelines. Thus, although the price indexes in the two periods are not directly comparable, the change in methodology involves all countries in the sample and its effects are captured by the good fixed effect. Second, the sample of countries and goods is significantly smaller than the cross-section analysis previously discussed. In fact, only two countries in the sample, Denmark and the United Kingdom, did not join the EMU.

Table 4: The Euro Shrunk the Band

	(All)	(3 Digit)	(4 Digit)
Euro	-0.639***	-0.711***	-0.499**
	(0.162)	(0.158)	(0.237)
R^2	0.126	0.167	0.044
# Observations	531	351	180

Results from OLS of equation (8). Robust std. error in parenthesis. Cluster: country pair. ***: significant at 99%, ** at 95%, * at 90%. Product fixed effects. All= all matched goods. 3 Digit= matched COICOP three-digit goods. 4 Digit= matched COICOP four-digit goods.

Table 4 shows the result. When a country pair started sharing the euro, their bands of inaction fell by approximately 64%. The magnitude of the effect is larger than the reduction of the bands of inaction experienced by Cyprus, Malta, Slovakia and Slovenia. The results are robust to changes in the sample of goods. Using the sample of three-digit COICOP goods, the euro effects is above 60%, while in the sample of four-digit COICOP goods, the entry in the EMU reduces bands of inaction by 50%. Running the cross section regression before and after the EMU creation highlights that the thresholds were somewhat larger between the EMU founders before entry.

I repeat (8) using the intercept and the slope of the time varying bands as dependent variables, and report the results in the online appendix. Contrary to the previous section, the initial threshold of the bands of inaction are declining after entry, which suggests that not only the EMU but other integration policies between the founders of the EMU have promoted economic integration reflected by the reduction in the bands of inaction. This result confirms the findings of Goldberg and Verboven (2005) and Rogers (2007) who documented price convergence prior to the creation of the EMU. As in the previous case, the results on the slope of the time varying bands are less robust. The coefficient is negative, but it attains statistical significance only in the sample of four-digit COICOP baskets of goods. This suggests that although part of the economic integration we observe between EMU founders is due to their historical connections and institutions, the EMU did facilitate a reduction in transaction costs afterwards. Prior to entry, the slope is not related to the future EMU membership.

5.3 Evidence from the Cross-Section

In this section, I explore the cross-country variation in bands of inaction to examine the relationship between euro and other transaction costs to relative PPP convergence. The objective is twofold. First, I can confirm the result of the previous section using the entire sample of countries available. Second, I can evaluate whether the threshold of the bands of inaction are related to several types of transaction costs as predicted by the model.

I estimate the TAR model using the entire sample of years and countries from Eurostat. I run the regression model outlined in (9) and report the results in table 5. The coefficient on the euro dummy is always negative and statistically significant. The interpretation of the coefficient is as follows. Since the average band of inaction is 0.118 and the coefficient on the Euro dummy in column (4) is -0.017, countries in the EMU have bands that are $\frac{0.017}{0.118}$ =14% smaller than the average band. Although such a result is not causal, it has a similar magnitude than the one shown in the previous section, which estimated a euro effect of 17%.

Table 5: The Euro and the Bands of Inaction

	(1)	(2)	(3)	(4)	(5)	(6)
Euro	-0.025***	-0.026***	-0.018***	-0.017***	-0.011***	-0.012***
	(0.005)	(0.005)	(0.006)	(0.006)	(0.004)	(0.004)
Log(Distance)		0.018***	0.018***	0.017***	0.013***	0.012***
		(0.005)	(0.005)	(0.004)	(0.003)	(0.003)
Common Border		0.007	0.006	0.007	0.003	0.004
		(0.008)	(0.008)	(0.008)	(0.005)	(0.006)
Common Language		-0.013*	-0.013*	-0.013*	-0.012**	-0.011*
		(0.008)	(0.007)	(0.007)	(0.005)	(0.006)
Island		0.035***	0.033***	0.029***	0.024***	0.026***
		(0.005)	(0.005)	(0.005)	(0.004)	(0.004)
Landlocked		0.008	0.006	0.003	-0.004	-0.001
		(0.005)	(0.006)	(0.006)	(0.004)	(0.005)
Fixed ERR			-0.011*	-0.014**	-0.022***	-0.016***
			(0.006)	(0.006)	(0.005)	(0.005)
Log(Expenditures)				-0.002*	-0.002***	-0.002***
				(0.001)	(0.001)	(0.001)
R^2	0.12	0.16	0.17	0.17	0.17	0.13
# Observations	9424	9424	9424	9424	8946	8946

Results from OLS of equation (9). Robust std. error in parenthesis. Cluster: country pair. ***: significant at 99%, ** at 95%, * at 90%. Product fixed effects. Sample: tradable four-digit COICOP goods. (5) WLS. (6) FGLS. Details in the main text.

Including geographical proxies for trade costs does not alter the euro coefficient, and bands of inaction reflect differences in trade costs across countries. The chosen geographical variables have the expected signs. More distant country pairs have wider bands of inaction, in line with the findings of Obstfeld and Taylor (1997) and of several studies documenting a positive relationship between deviations from LOP and distance (Engel and Rogers, 1996).[31] Bands of inaction decline if two countries share a common language while they increase if one of the countries is an island. The results are in line with the findings of Lutz (2004), who document that price dispersion across countries increases with distance and falls with commonality of language.

The currency union effect on the bands of inaction is more than that arising from a simple fixed exchange rate. In fact, the coefficient on the euro dummy variable is negative and statistically significant when I control for the presence of a fixed exchange rate regime, despite declining somewhat in magnitude. This result is consistent with the findings of Cavallo et al. (2014), who find that price deviations from the LOP are smaller in currency unions relative to countries that have pegged currencies.

The coefficient on Expenditures is negative and statistically significant. The larger the size of the country pair is, the smaller are the bands of inaction. From the simple theoretical model, such a result supports the hypothesis that arbitrageurs must pay the fixed costs, not only the variable costs of arbitrage. The result is in contrast with Lutz (2004) who finds that price dispersion increases with market size.

A potential concern may arise from the fact that the dependent variable of (9) is the outcome of an estimation algorithm and, thus, carry a standard error. The results shown thus far implicitly assume that the precision with which the thresholds of the bands of inaction are estimated is constant across goods and country pair. To address the concern, I follow the procedures described by Lewis and Linzer (2005) in the context of Estimated Dependent Variables. First, I consider Weighted Least Squares (WLS) where the weights are the inverse of the standard error of the bands of inaction. Second, I apply Feasible Generalized Least Squares (FGLS), which assumes that the weights are a function of the inverse of the standard error of the bands of inaction. The results are in column (6) and (7) and are virtually identical to the baseline of column (5). Details on WLS and FGLS are in the online appendix.

The effect of the euro on the bands of inaction are quite heterogeneous across goods. I estimate (9) for each basket of goods separately and report the full set results in the online appendix. The largest drop in the bands of inaction associated with the EMU is registered by printed matter, garments, and meat. For these baskets, country pairs within the EMU have approximately 40% smaller bands than the average basket-specific band. For a few baskets of goods, such as fruit and cars, belonging to the EMU is associated with wider bands of inaction.

The effects of the other explanatory variables are also heterogeneous across goods. Distance, being an island, and fixed exchange rate regime have the most consistent effects across goods, in

[31] Engel and Rogers (1996) also document a border effect that drives deviations from LOP across countries, but Gorodnichenko and Tesar (2009) show that such border effect is simply the result of trade between countries with different distributions of prices.

line with the baseline result. Landlocked countries have higher thresholds of the bands of inaction for vegetables, garments, and games. Finally, examining the coefficient on Expenditures suggests relatively high fixed costs of transaction for solid fuels and motor cycles.

Threshold with a Linear Trend. Similarly to the results from the previous section, the relationship between the slope of the time varying bands of inaction and the euro is less robust. Using the coefficient on the trend b_{ij}^{g} as dependent variable, I estimate the regression model outlined in (9) and report the results in the online appendix. The coefficient on the euro dummy is negative, but statistically insignificant. The lack of statistical significance is driven by few outlier country pairs, as the estimation on the euro coefficient becomes more efficient as I restrict the sample of country pairs to members of the EU, Schengen Area, and fixed exchange rate regimes countries, anticipating a few of the robustness checks I outline here after.[32]

Furthermore, there is high heterogeneity in results across baskets of goods. In fact, the euro is correlated with a narrowing of the bands for carpets, clothing, garments, solid fuels, plants and flowers, durables for recreation, and photographic and information processing equipment. In contrast, the euro is associated with a widening of the thresholds for bread and cereals, fish, sugar, coffee and tea, spirits, beer, and jewellery. Details are in the online appendix.

5.3.1 Robustness

Unlike the study of entry to the EMU, the thresholds used in this section represent the average threshold from 1999 to 2016. Hence, considering the case of new entrants, the threshold represent the average between the higher threshold prior to entry, and the lower threshold post-entry. To capture this feature of the estimation, I count the number of years in which a given country pair share the same currency, and then repeat (9) using the number of years as the explanatory variable. I find a negative and statistically significant coefficient: the longer the pair has been in the common currency, the stronger is the associated reduction in the band of inaction. In an alternative approach, I create three new dummy variables to capture three possible relationships: 1) a pair formed by founders of the euro, 2) a pair formed by new members of the EMU, and 3) a pair formed by a founder and a new member. The estimated coefficients on the three dummies are negative and statistically significant for all but the pairs involving two new members. Such a result might be due to new entrants that are in the sample for a limited number of years.[33]

The coefficient on the euro dummy might be biased by the concurrent belonging of EMU countries to the EU and of some of them to the Schengen Area. To address this issue, I restrict the

[32] I also estimate the regression model using the intercept of the time varying bands as dependent variable finding results similar to using the average bands of inaction. As the two variables are highly correlated, the result is not surprising.

[33] As previously mentioned, the historical stronger integration between the founders of the EMU might drive the results. However, the results are robust to dropping all country pairs in which both countries are founders of the euro.

sample of country pairs used in regression (9), conditioning on those pairs that belong to the EU, the Schengen Area, or both. The effect of the EMU persists: the coefficient on the euro is negative and statistically significant in all specifications. Conditional on the country pairs belonging to the EU, sharing the same currency reduces the bands of inaction by 12% relative to the average band in the EU. Country pairs in the Schengen Area and EMU have bands of inaction that are 26% smaller than the average band in the Schengen Area. Moreover, countries in the EMU have bands of inaction that are 23% smaller than the average band in the EU and Schengen Area combined.[34]

Finally, results are robust to using alternative measures of exchange rate volatility. The negative coefficient on the euro persists controlling for the variance of bilateral exchange rates. Moreover, results are robust to restricting the sample to country pairs in a fixed regime and in the EU. The currency union effect is robust to alternative measures of trade costs, using a non linear function of the log of distance (Simonovska, 2015) and the indicators of distance region from Eaton and Kortum (2002).

Variable and Fixed Costs. Although the empirical specification does not allow for a distinction between fixed and variable costs of arbitrage, I examine the presence of fixed costs of arbitrage following two strategies. First, I use a measure for the ease with which businesses are created as a proxy for the fixed transaction costs associated with arbitrage. To proxy for the ease with which business are created in a country, I use the average placement of such country for ease of doing business according to the World Bank for the years available (2015 and 2016). The average placement can range between 1 (best country for opening businesses) to 190 (worst country). In regression (9), I control for the minimum value of the ease of doing business between the two countries in the pair ij. I choose the minimum value between the two countries because it is the minimum fixed cost of arbitrage that defines the thresholds of the bands of inaction. The variable is a proxy for all frictions that may be correlated with the fixed cost of arbitrage. I find that the easier it is to do business, the smaller the bands of inaction.

Second, the presence of fixed costs of arbitrage creates a negative relationship between the size of the market and the bands of inaction. Such a negative relationship is robust to changes in the measures of expenditures used in the regression, such as total household consumption and GDP weighted by consumption shares. Moreover, not only the size of the market reduces the bands of inaction, but also the size of the distribution network improves price convergence. Using data from Eurostat, I find that the larger the turnover of retailers and wholesalers, the smaller the bands of inaction. Finally, the rise in Internet use facilitates arbitrage opportunities between traditional retailers and online stores. I proxy for the size of online markets using Eurostat data on the share of individuals in one country having made an online purchase in the previous three months. The

[34] An even stronger test is to condition on countries that are members of the EU, that are members of the Schengen Area, and that share a common border. An example is to compare the bands of inaction between Austria, Czech Republic, and Slovakia. Even in this case, the coefficient on the euro is negative and statistically significant. Details in online appendix.

result highlights a negative relationship between bands of inaction and online markets.

Non-Tradable Goods. A test for the sensibility of the empirical model is to examine whether transaction costs and a common currency affect the bands of inaction for non-tradable goods. I follow the same procedure described for tradable goods and estimate the bands of inaction for 36 non-tradable four-digit COICOP goods. The average threshold for the bands of inaction is $\pm 13.8\%$, which is larger than the average across tradable goods.[35] Moreover, the results from the regression (9) on the sample of non-tradable goods are as expected: the euro has an insignificant impact on their bands of inaction. In addition, other variables that play a role in explaining the bands for tradable goods, such as distance and the EU, do not have a significant effect for non-tradable goods.

The bands of inaction of non-tradable goods decrease with the size of the two countries. Moreover, the bands are lower when a country pair shares a common language while they are higher when one of the countries in the pair is an island. We can interpret the result in light of the findings of Engel and Rogers (2004), who document that in European countries, price dispersion of both tradable and non-tradable goods fell in the 1990s. The authors argue that the integration of factors' markets drives the price convergence in the non-tradable sector. The results shown here suggest that the size of the country pair, the commonality of language, and the country's isolation (captured by the island dummy) are contributing factors to the integration in the market for factors of production.

Speed of Convergence and Price Dispersion. Does the euro increase the speed of convergence of relative prices? To answer this question, I test whether the euro affects the estimated λ_{ij}^g, the autoregressive coefficient outside the band of inaction (7). In fact, the larger the absolute value of λ_{ij}^g is, the faster is the convergence of relative prices toward the band. I regress λ_{ij}^g on the euro dummy variable and the controls used in the empirical model (9). The regression results indicate that the euro is associated with a larger absolute value of λ_{ij}^g. The common currency is not only related to lower bands of inaction but also with faster convergence when prices are outside the band. The result is in line with Bergin et al. (2017), who find that the speed of convergence to PPP is higher in the EMU.

Another related question is whether the euro reduces the volatility of the demeaned log PPP deviations x_{ijt}^g (6). To examine the relationship between the volatility of prices over time and a common currency, I regress the standard deviation of x_{ijt}^g on the same controls used in (5), finding

[35] The result is consistent with Crucini et al. (2005), who document that price dispersion across destinations declines with the tradability of a good. The relative small difference of the average band between tradable and non-tradable goods is surprising. Such a result may suggest the presence of a particularly strong aggregation bias that underestimates the bands of inaction for non-tradable good. Alternatively, a possible concern may be that the classification of goods by Sturm et al. (2009) is not appropriate. As robustness, I consider the effects of the euro on the bands of inaction for all goods, both tradable and non-tradable. The effect of the euro is robust to this change in the sample of goods.

a negative and statistically significant coefficient on the euro dummy. The result is similar in spirit to the findings of Cavallo et al. (2015), who document a reduction in the price dispersion across goods between members of a currency union. This result is not surprising. If the euro reduces the bands of inaction, the volatility of prices must fall.

Aggregate Price Indexes. Finally, the results are robust to changing the set of goods employed for the analysis. I repeat the empirical exercise using three-digit COICOP goods, finding that the euro reduces the bands of inaction relative to the average by 9%. The descriptive results and the coefficients on the other controls employed are remarkably similar to the baseline case. The finding suggests that neither the aggregation bias, not the negative relationship between number of goods in the basket and bands of inaction, drive the main result of the study. The online appendix shows the detailed results.

6 Conclusions

I have argued that a currency union reduces the bands of inaction of relative price indexes via a drop in transaction costs. The results are in line with those of Lutz (2004), Allington et al. (2005), Cavallo et al. (2015), and Cavallo et al. (2014), who document that price dispersion declines within a currency union. Rather than focusing on deviations of prices from the LOP to assess the integration arising in the EMU, this study used a TAR model (Obstfeld and Taylor, 1997) to estimate the bands of inaction. I examine the entry of Cyprus, Malta, Slovakia, and Slovenia to the EMU, finding that their bands of inaction fell by 17% after entry. Results from a cross-section analysis further confirm the results. The EMU eliminates costs associated with hedging of exchange rate risk and with currency exchange (Anderson and van Wincoop, 2003). Moreover, the EMU reduces information frictions (Boivin et al., 2012), as it improves the ease with which consumers compare prices across locations (Mussweiler and Strack, 2004).

Although the study focused on one particular channel — transaction costs — for why prices in a currency union tend to converge, it does not exclude other possible mechanisms. In particular, Cavallo et al. (2014) document that the price of a good at its introduction is a major component of the good-specific real exchange rate. This suggests that firms could internalize arbitrage possibilities and thereby set prices within a smaller range to prevent arbitrage. Other channels are worth investigating, such as pricing to currency area or, as argued by (Cavallo et al., 2014), charging the same prices so as not to anger firms' customers (Rotemberg, 2005).

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7 Appendix

7.1 Goods and Country Details

Table 6 shows the list of countries and the date of entry in European Institutions. Iceland, Norway, and Switzerland belong to the European Free Trade Association (EFTA).[36] In addition, Iceland and Norway belong to the European Economic Area which enables them to participate in the EU's internal market.[37] Switzerland has a free trade agreement with the EU signed in 1972, extended in 1999 to trade in agricultural products, and again extended in 2004 to processed agricultural products.[38] Finally, Turkey has been in a customs union with the EU since 1996 for all goods but agricultural products, coal, and steel.[39]. Table 7 divides the four-digit COICOP baskets of goods used in the empirical analysis in tradable and non-tradable according to Sturm et al. (2009).

7.2 Estimating the Bands of Inaction

I employ a grid search to maximize the likelihood of the TAR model against a null AR(1) model. Let $L_n(\lambda, \sigma)$ be the log likelihood function of the null AR(1) model, defined as follows:

$$\Delta x^g_{ijt} = \lambda^g_{ij} x^g_{ji\,t-1} + e^g_{ijt} \tag{10}$$

where $e^g_{jit} \sim N(0, \sigma^{2g})$. The likelihood of the null model is given by

$$L^g_{n,ij}(\lambda^g_{ij}, \sigma^g_{ij}) = -\sum_t \frac{1}{2}\left(\log(2\pi) + \log(\sigma^{g2}_{ij}) + \frac{e^{g2}_{ijt}}{\sigma^{g2}_{ij}}\right)$$

Similarly, let $L^g_{TAR,ij}(\lambda^g_{ij}, \sigma^{g,out}_{ij}, \sigma^{g,in}, c^g_{ij})$ be the log likelihood function of the TAR model (7):

$$\begin{aligned} L^g_{TAR,ij}(\lambda^g_{ij}, \sigma^{g,out}_{ij}, \sigma^{g,in}_{ij}, c^g_{ij}) = & -\sum_{I_{in}(x^g_{ij\,t-1})=1} \frac{1}{2}\left(\log(2\pi) + \log(\sigma^{g,in2}_{ij}) + \frac{e^{g,in2}_{ijt}}{\sigma^{g,in2}_{ij}}\right) \\ & -\sum_{I_{out}(x^g_{ij\,t-1})=1} \frac{1}{2}\left(\log(2\pi) + \log(\sigma^{g,out2}_{ij})) + \frac{e^{g,out2}_{ijt}}{\sigma^{g,out2}_{ij}}\right) \end{aligned} \tag{11}$$

where $I_{in}(x^g_{ij\,t-1}) = 1(|x^g_{ij\,t-1}| \leq c^g_{jit})$ and $I_{out}(x^g_{ij\,t-1}) = 1(|x^g_{ij\,t-1}| > c^g_{jit})$ are indicator functions.

Let us first consider the costant threshold case. I estimate the parameters of the model by a grid search on c_{ji}, which maximizes the log likelihood ratio $LLR^g_{ij} = 2(L^g_{TAR,ij} - L^g_{n,ij})$. The procedure

[36] http://www.efta.int. Iceland was a candidate to join the EU from 2010 but negotiations froze in 2013. Sweden, Austria, and Finland were in the EFTA until 1995, when they joined the EU.

[37] http://www.efta.int/eea/eea-agreement

[38] http://eeas.europa.eu/switzerland/index_en.htm

[39] http://ec.europa.eu/taxation_customs/customs/customs_duties/rules_origin/customs_unions/article_414_en.htm

Table 6: List of Countries and Date of Access to EU, EMU, and Schengen Area

	EU (or EEC)	EMU	Schengen Area
Austria	Jan 1995	Jan 1999	Dec 1997
Belgium	Founder	Jan 1999	March 1995
Bulgaria	Jan 2007		
Croatia	Jul 2013		
Cyprus	May 2004	Jan 2008	
Czech Republic	May 2004		Dec 2007
Denmark	Jan 1973		Mar 2001
Estonia	May 2004	Jan 2011	Dec 2007
Finland	Jan 1995	Jan 1999	Mar 2001
France	Founder	Jan 1999	Mar 1995
Germany	Founder	Jan 1999	Mar 1995
Greece	Jan 1981	Jan 2001	Jan 2000
Hungary	May 2004		Dec 2007
Iceland			Mar 2001
Ireland	Jan 1973	Jan 1999	
Italy	Founder	Jan 1999	Oct 1997
Latvia	May 2004	Jan 2014	Dec 2007
Lithuania	May 2004		Dec 2007
Luxembourg	Founder	Jan 1999	Mar 1995
Malta	May 2004	Jan 2008	Dec 2007
Netherlands	Founder	Jan 1999	Mar 1995
Norway			Mar 2001
Poland	May 2004		Dec 2007
Portugal	Jan 1986	Jan 1999	Mar 1995
Romania	Jan 2007		
Slovakia	May 2004	Jan 2009	Dec 2007
Slovenia	May 2004	Jan 2007	Dec 2007
Spain	Jan 1986	Jan 1999	Mar 1995
Sweden	Jan 1995		Mar 2001
Switzerland			Dec 2008
Turkey			
United Kingdom	Jan 1973		

is as follows. First, I find the 10^{th} and 90^{th} percentiles of $|x_{jit}|$ and then divide the interval in steps of 0.001 width. Such steps are the candidate solutions for c^g_{ji}. For each possible c^g_{ji}, I partition the sample of $x^g_{ij\,t-1}$ into observations within or outside the band. Then, I use ordinary least squares (OLS) to compute the other parameters of the model (λ^g_{ij}, $\sigma^{g,out}_{ij}$ and $\sigma^{g,in}_{ij}$) and the log likelihood previously described. Finally, I pick the c^g_{ji} that maximizes the log likelihood ratio.

To estimate the standard errors for the thresholds of the band of inaction, I follow Caner and Hansen (2001) and use a bootstrap method. Let us focus on a given basket g for the country pair ij. Let the estimated threshold be $\hat{c}^g_{ij}$ and the estimated autoregressive coefficient outside of the bands be $\hat{\lambda}^g_{ij}$. Let T denote the total number of months available in the data for the good and country pair. Given $\hat{c}^g_{ij}$, $\hat{\lambda}^g_{ij}$, and the data on relative price indexes changes Δx^g_{ijt} and levels $x^g_{ij\,t}$, I compute the empirical distribution of the residuals $\hat{e}^{g,in}_{ijt}$ and $\hat{e}^{g,out}_{ijt}$ as follows:

$$
\begin{aligned}
\hat{e}^{g,out}_{ijt} &= \Delta x^g_{ijt} - \lambda^g_{ij}(x^g_{ij\,t-1} - c^g_{ij}) && \text{if } x^g_{ij\,t-1} > c^g_{ij} \\
\hat{e}^{g,in}_{ijt} &= \Delta x^g_{ijt} && \text{if } c^g_{ij} \geq x^g_{ij\,t-1} \geq -c^g_{ij} \\
\hat{e}^{g,out}_{ijt} &= \Delta x^g_{ijt} - \lambda^g_{ij}(x^g_{ij\,t-1} + c^g_{ij}) && \text{if } -c^g_{ij} > x^g_{ij\,t-1}
\end{aligned}
\tag{12}
$$

Let $F^{in} = \{\hat{e}^{g,in}_{ijt}\}$ and $F^{out} = \{\hat{e}^{g,out}_{ijt}\}$ denote the sets containing the empirical estimates of the residuals inside and outside the bands of inaction.

I simulate the bootstrap evolution of relative price indexes $\tilde{x}^g_{ij\,t}$ using the TAR model (7), for $t = 1, ..., T$. The thresholds of the band are the estimated thresholds $\hat{c}^g_{ij}$, the autoregressive coefficient is the estimated one $\hat{\lambda}^g_{ij}$, and the shocks $e^{g,out}_{ijt}$ and $e^{g,in}_{ijt}$ are drawn from F^{out} and F^{in}. I choose as the initial value for the bootstrap simulation $\tilde{x}^g_{ij\,t=0} = 0$, namely that relative PPP holds. Using the simulated time series, I estimate the threshold of the band of inaction $\tilde{c}^g_{ij}$, using the algorithm described in the previous section. I repeat the procedure 100 times for each basket and country pair, and compute the standard error of the bootstrap-estimated thresholds $\tilde{c}^g_{ij}$.

Since the algorithm that estimates the thresholds uses a grid search over the sample values of relative price indexes, for the bootstrap algorithm to produce results in a reasonable amount of time, the sample values of the simulated relative price indexes has to be somehow contained. For this reason, I only compute the standard errors for the goods and country pairs with an estimated autoregressive coefficient that satisfies the following inequality: $-1 < \hat{\lambda}^g_{ij} < 0$.

Let us now consider the case in which the thresholds have a linear trend:

$$c^g_{ijt} = a^g_{ij} + b^g_{ij}(t-1) \qquad t = 1, ..., T$$

To estimate the two parameters a^g_{ij} and b^g_{ij} I proceed as follows. First, at $t = 1$, $c^g_{ij1} = a^g_{ij}$. Second, I assume that the average c^g_{ijt} across time equals the threshold estimated in the baseline model with constant thresholds: $\sum_{t=1}^{T} T^{-1} c^g_{ijt} = c^g_{ij}$. Due to this simplifying assumption, $b^g_{ij} = \frac{T(c^g_{ij} - c^g_{ij1})}{\sum_{t=1}^{T}(t-1)}$.

Therefore, the threshold at time t is given by:

$$c^g_{ijt} = c^g_{ij1} + \frac{T(c^g_{ij} - c^g_{ij1})}{\sum_{t=1}^{T}(t-1)}(t-1) \qquad t = 1, ..., T \tag{13}$$

Given the estimated band of inaction in the baseline model c^g_{ij}, I estimate the new set of parameters via maximum likelihood. In this case, the grid search is done over the initial level of the threshold c^g_{ij1}. The procedure is as follows. First, I find the 10^{th} and 90^{th} percentiles of $|x_{ijt}|$ and then divide the interval in steps of 0.001 width. Such steps are the candidate solutions for c^g_{ij1}. For each possible c^g_{ij1}, I partition the sample of $x^g_{ij\,t-1}$ into observations within or outside the band. Then, I use OLS to compute the other parameters of the model (λ^g_{ij}, $\sigma^{g,out}_{ij}$ and $\sigma^{g,in}_{ij}$) and the log likelihood previously described. Finally, I pick the c^g_{ij1} that maximizes the log likelihood ratio.

Table 7: Tradable and Non-Tradable Four-Digit COICOP

Tradable		Non tradable	
Code	Description	Code	Description
CP0111	Bread and cereals (B)	CP0314	Cleaning of clothing (B)
CP0112	Meat (B)	CP0432	Services for repair of dwelling
CP0113	Fish (B)	CP0441	Water supply
CP0114	Dairy products (B)	CP0442	Refuse collection
CP0115	Oils and fats (B)	CP0443	Sewerage
CP0116	Fruit (B)	CP0444	Services to dwelling
CP0117	Vegerables (B)	CP0451	Electricity (B)
CP0118	Sugar	CP0452	Gas
CP0119	Food n.e.c. (B)	CP0455	Heat energy
CP0121	Coffe and tea (B)	CP0513	Repair of furniture
CP0122	Soft drinks (B)	CP0533	Repair of household appliances
CP0211	Spirits (B)	CP0562	Household services
CP0212	Wine (B)	CP0621-0623	Medical services
CP0213	Beer	CP0622	Dental services
CP0311	Clothing materials	CP0723	Maintenance of transport (B)
CP0312	Garments (B)	CP0724	Services to transport (B)
CP0313	Clothing n.e.c. (B)	CP0731	Transport by railway
CP0431	Materials for repair of dwelling	CP0732	Transport by road (B)
CP0453	Liquid fuels	CP0733	Transport by air (B)
CP0454	Solid fuels	CP0734	Transport by sea
CP0511	Furniture	CP0735	Combined transport
CP0512	Carpets	CP0736	Other transport services
CP0531-0532	Household appliances	CP0830	Telephone services
CP0561	Non-durable household goods	CP0915	Repair of equipment
CP0611	Pharmaceutical (B)	CP0923	Maintenance of durables for recreation
CP0612-0613	Medical products	CP0941	Recreational services (B)
CP0711	Cars	CP0942	Cultural services (B)
CP0712-0714	Motor cycles	CP0951	Books
CP0721	Accessories for transport	CP0952	Newspapers
CP0722	Fuels for transport	CP1111	Restaurants (B)
CP0911	Equipment for sound recording	CP1112	Canteens
CP0912	Photographic equipment	CP1211	Hairdressing (B)
CP0913	Information processing eq. (B)	CP1252	Insurance for dwelling
CP0914	Recording media	CP1253	Insurance for health
CP0921-0922	Durables for recreation	CP1254	Insurance for transport
CP0931	Games (B)	CP1255	Other insurance
CP0932	Equipment for sport (B)		
CP0933	Plants and flowers		
CP0934-0935	Pets		
CP0953-0954	Printed matter		
CP1212-1213	Electrical appliances (B)		
CP1231	Jewellery		
CP1232	Personal effects		

There are 36 non-tradable goods. Telephone and telefax equipment (COICOP 08.20) is dropped because it contains both a tradable (purchases) and a non-tradable (repair) component. (B) indicates whether the basket is included in the regressions that use a balanced panel of commodity points.