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ABSTRACT

Did Inflation Really Soar After the Euro Cash Changeover? Indirect Evidence from ATM Withdrawals

The introduction of the euro notes and coins during the first months of 2002 was followed by a lively debate on the alleged inflationary effects of the new currency. In Italy, as in the rest of the euro area, survey-based measures signaled a much sharper rise in inflation than measured by the official price indices, whose quality was called into question. In this paper we gather indirect evidence on the behaviour of prices from the analysis of cash withdrawals from ATM and their determinants. Since these data do not rely on official inflation statistics, they provide an independent check for the latter. We present a simple theoretical model in which the relationship between aggregate ATM withdrawals and aggregate expenditure is not homogenous of degree one in the price level, a prediction which is strongly supported by the data. This feature allows us to test the hypothesis that, after the introduction of the euro notes and coins, consumer prices underwent an increase not recorded by official inflation statistics. We do not find evidence in support of this hypothesis.

JEL Classification: E31 and E41
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"Whatever the experts say, many European consumers still feel retailers are masking price increases with the changeover to the euro" (Wall Street Journal Europe, 28.01.02)

"Two out of three eurozone consumers felt they were ripped off by retailers during the changeover to pricing in euros, according to the European Commission…. Germany, France and Netherlands were the countries with the highest percentage of people feeling cheated…" (Financial Times, 01.03.02)

"German consumers dubbed the currency the Teuro (teuer is German for expensive). […] Some consumers believe higher prices were the result of retailers rounding up prices as they switched out of their old national currencies into the euro. However, EU statisticians insisted prices had not been affected" (Financial Times, 12.12.02)

1. Introduction

There is a widespread perception among the citizens of the Euro area that the introduction of the euro notes and coins in the first months of 2002 spurred a rise in inflation that was much sharper than measured by the national statistical offices (see European Central Bank, 2002a, 2003a,b). This phenomenon, illustrated in Figure 1, has been the subject of countless newspaper articles and of several official speeches by policymakers and politicians.[1] While the quotations reported above refer to 2002, the perception that the euro brought about higher inflation is still vivid at present.[2]

It is somewhat puzzling that a change in the unit of account might have an impact on the inflation rate, and indeed, a few hypotheses have been set forth to explain why perceptions of high inflation are not necessarily at odds with official statistics (see Del Giovane and Sabbatini (2004); Traut-Mattausch *et al*, 2004). Still, these explanations, which emphasize the role of psychological factors, are hard to test. Therefore, the nagging question remains: have the official statistics failed to capture large prices increases? Clearly, a positive answer would damage the credibility of monetary policy, which is based on official inflation statistics, and could eventually be engraved in consumers'

[1] In May 2002 Prof. O. Issing gave a speech in Mainz on "Der Euro - eine stabile Währung". After the speech, the first question from the audience was about the "teuro" phenomenon. Seeing the look of disbelief with which his explanation was met, Issing replied: "You seem not to believe me. And even my wife doesn't believe me". This sentence found wide coverage in the German press.

[2] More than two years after the changeover, President Trichet still deemed it necessary to reassure European customers on this issue: "European citizens who still perceive that inflation is higher than measured by official indices should be

views over European institutions. But even lack of a clear answer to this question can have negative consequences on credibility, to the extent that the actors involved persist in their positions – the policymakers believing in the goodness of the official price measures, the consumers believing the opposite.

The main obstacle faced by researchers in tackling this issue is the absence of reliable alternative inflation measures. The thesis that price increases have been far larger than measured by the national statistical offices, reflected in the measures of perceived inflation, is mostly based on anecdotal evidence, or on the prices of a few selected goods; at best, on price surveys that cannot be compared with those underlying official indexes, based on an exhaustive list of goods and services and encompassing in terms of geographical coverage.

This paper investigates the dynamics of the price level in Italy after the introduction of the euro notes and coins (the so-called cash changeover) by using data on currency demand patterns. We believe that this inference, albeit indirect, is useful because it relies on data that are collected and assembled by central banks, with methodologies that are completely independent from those used by the national statistical offices. The basic steps of our investigation can be summarized as follows. We setup a simple theoretical model of ATM withdrawals and use it as a guide for our econometric analysis of the determinants of ATM withdrawals *prior* to the changeover, when official statistics were arguably correct. The model suggests that the relationship between aggregate ATM withdrawals and aggregate expenditure is not homogenous in the price level. This feature, which finds strong support in the data, implies that price level dynamics can be deduced from the observed nominal time series for withdrawals and expenditure. It is shown that if official data on prices are biased after the changeover, but data on withdrawals and expenditure are not, extending the estimation period to 2002-03 should cause instability in the estimated coefficients, which can be captured econometrically. Formally, we test the null hypothesis that the increase in consumer prices is correctly measured by official statistics after the changeover. Both a price-level bias and an inflation-bias hypothesis are formulated and tested. In a nutshell, we fail to find evidence consistent with the occurrence of a price hike after the changeover. A counterfactual exercise, based on simulated data, suggests that this result cannot be ascribed to a low power of the statistical test.

Several reasons warrant our focus on Italy. First, this country is broadly representative of the euro area in terms of the discrepancy between official and perceived inflation (fig. 1). Also, quarterly

assured that the official measures are accurate and that we will continue to maintain price stability in the future" (introductory statement after the Governing council meeting of April 2004).

data on cash withdrawals are available, whereas for other euro area countries, to our knowledge, comparable data are available only at an annual frequency. Several reasons also warrant our focus on the flow of currency withdrawn from the ATM circuit, rather than on more traditional monetary aggregates. The stock of currency experienced a strong decline from the beginning of 2001, apparently reflecting weak demand of banknotes as a store of wealth due to the approaching currency changeover (see e.g. European Central Bank, [2002b]). Among the traditional monetary aggregates, M1 is strongly affected by the erratic behavior of currency. M2 and M3 are typically less related to transactions; in addition, over the recent past their dynamics has been strongly influenced by portfolio reasons, as repeatedly argued by the European Central Bank. By contrast, there is no obvious reason why cash demanded for transactions – undoubtedly the main determinant of ATM withdrawals – should have been affected by these same factors.

Related studies on the price consequences of the changeover have mainly concentrated on industries where the effects of the changeover were allegedly strong. Hobijn *et al.* (2004) and Gaiotti and Lippi (2004) analyze the dynamics of restaurant prices, finding some evidence consistent with the price-hike hypothesis. Such evidence, however, is rooted in industry-specific features (e.g. large menu costs) and cannot be used as a proxy for the dynamics of the general price level. An approach more closely related to ours is in Marini *et al.* (2004). They infer inflation dynamics in Italy from a consistency check of households' median real income and (self-assessed) financial situation based on survey data, concluding that after the changeover official statistics underestimated inflation by about 6 percentage points. The survey-based nature of the data used by these authors, however, suggests that this finding is best interpreted as a measure of perceived, rather than actual inflation, an interpretation which is consistent with the evidence of Figure 1.

The paper is organized as follows. In the next Section we present a model of demand for ATM withdrawals which is used as a guideline in the empirical analysis of Section 3. A final Section summarizes the main findings.

2. A simple model of ATM withdrawals

This section presents a simple model of ATM cash withdrawals which provides a useful guidance for the empirical analysis of Section 3. The model explains aggregate (i.e. nationwide) withdrawals, since the data available to us are aggregated. It does so by focusing on an "intensive" and an "extensive" margin. The former concerns the relative intensity with which agents who possess an ATM card use cash relative to substitute means of payments. The latter pertains to the fraction of citizen who possess an ATM card, which is important in explaining aggregate dynamics given the strong diffusion of ATM cards over the sample period.[3] We consider the two segments one at the time.

Before proceeding, it is worth noting that attention to the theoretical underpinnings of the analysis may be warranted by a few differences between the present context and a traditional money demand equation. While in the latter the dependent variable is a stock, we focus here on the currency flow (withdrawn from ATM) used to pay for consumption expenditure. Partly as a consequence of this, the explanatory variables of our demand equation need not be the same as those used in a standard money demand. Here the relevant margin is between the costs of alternative means of payments (e.g. cash versus credit) while in the case of money demand it is between the yields of alternative assets (e.g. money versus non monetary assets).[4] Additional explanatory variables are also required by our modeling of the "extensive" margin, typically overlooked in a money demand function.

Let i index an agent who possesses an ATM card, and let E^i denote her (exogenous) nominal consumption expenditure over a given time span (e.g. a quarter). To pay for E^i the agent can use cash, C^i, assumed to be withdrawn from ATMs, or alternative payment instruments, which we denote by Q^i, e.g. credit cards, points of sale (POS), transactions performed using a debit card, cheques etc.[5] Both C^i and Q^i are flows, so that over the period $C^i + Q^i = E^i$. The agent's choice concerns the proportion $k^i = C^i/E^i$ of expenditure to be financed with cash. We now turn to the agent's choice about the optimal proportion of cash expenditure k^i.

When deciding whether to use cash or "substitutes" to carry out transactions, the agent is faced with the following problem. The cost of cash relative to "substitutes" (per unit of expenditure) is R, which can be thought of as the nominal interest rate. This captures the fact that when nominal interest rates are higher, cash holdings give rise to a greater amount of forgone interest on deposits. Using a substitute payment instrument, such as a credit card, allows the agents to avoid bearing such cost. Moreover, it is assumed that ATM withdrawals succeed with probability $z(d^{ATM}, k^i)$. The agent may be unable to withdraw because she cannot find a conveniently located ATM, or due to network downtime. Thus, this probability is assumed to be increasing in the parameter d^{ATM}, measuring ATM diffusion over the national territory. We also postulate that the more the agent relies on ATM withdrawals, the more likely she is to run into a malfunctioning. Thus, $z(d^{ATM}, k^i)$ has $z_1>0$ and $z_2<0$. With probability

3 In Italy the proportion of households owning an ATM card rose from 15 per cent in 1989 to 55 per cent in 2002.

4 We shall see below that while the traditional demand for currency is a decreasing function of the nominal rate of interest, this is not necessarily true in our model.

5 The assumption that all cash is withdrawn from the ATM can be easily relaxed (e.g. assuming that some cash is also withdrawn from the bank desk) without altering the qualitative implications of the model.

$(1–z)$ the agent fails to withdraw and bears a cost f, $f>R$, which can be thought of as the cost of time lost searching for another cash dispenser. Since the cost of time is presumably greater for wealthier people, f is assumed to be an increasing function of individual wealth (proxied by real consumption expenditure). Thus $f\left(E^i/P\right)$ has $f_1>0$ (P denotes the price level).

The cost of using cash substitutes is normalized to zero (recall that R measures the relative cost of cash vs. substitutes). By symmetry with C^i, payments settled with non-cash instruments Q^i succeed with probability $s(d^Q,1–k^i)$, where d^Q measures the diffusion of the network accepting non-cash means of payments (e.g. the POS network) and $(1–k^i)$ denotes the proportion of expenditure settled with Q^i, and $s_1>0$, $s_2<0$. With probability $(1–s)$ the agent is unable to resort to Q^i and incurs the cost f.[6]

Overlooking i superscripts, the optimal choice of k^i solves the following problem:

$$\min_{k} \; k\,[R\,z(d^{ATM},k)+(1-z(d^{ATM},k))f] + (1-k)\,[1-s(d^Q,1-k)]f \tag{1}$$

where the terms multiplied by k and by $(1–k)$ denote, respectively, the expected cost of "cash expenditure" and of "non-cash" expenditure. The first order condition for an interior minimum is:[7]

$$(R-f)(z+kz_2)+f(s+(1-k)s_2)=0 \tag{2}$$

Using equation (2) we are able to characterize the optimal proportion of cash expenditure as an implicit function: $k^i = k^i(d^{ATM}, d^Q, \frac{E^i}{P}, R)$. The optimal amount of cash expenditure for a typical holder of an ATM card is thus determined by:

$$C^i = k^i(d^{ATM}, d^Q, \frac{E^i}{P}, R)\,E^i \tag{3}$$

Let us momentarily postpone the discussion of equation (3) and consider the aggregation problem. Let n be the number of agents with an ATM card and N be the population size. Denote aggregate withdrawals and expenditure for the group of ATM cardholders by, respectively, $C=nC^i$ and $E=nE^i$. Further, let $q \equiv E/\overline{E}$, where $\overline{E}$ denotes aggregate expenditure. In order to bring equation (3) to

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6 To keep matters simple we assume that the agent incurs the cost f when either an ATM withdrawal or a non-cash payment fails, ruling out the possibility of a sequential substitution between the two means of payments (e.g. to try to pay with credit if the ATM withdrawal fails, and *viceversa*).

7 The second order condition is $(R-f)[2z_2+kz_{22}]-f[2s_2+(1-k)s_{22}]>0$. Sufficient conditions for the second order condition to hold are $z_{22}<0$ and $s_{22}<0$.

the data, we model q as $q = m\dfrac{n}{N}$.[8] Thus, $E^i = \dfrac{m}{N}\overline{E}$. Substituting this expression in (3) and multiplying both sides by n yields an expression relating the aggregate flow of ATM withdrawals to aggregate consumption expenditure, both expressed in real terms, as:

$$(4) \qquad \frac{C}{P} = k\left(d^{ATM}, d^{Q}, \frac{m}{N}\frac{\overline{E}}{P}, R\right)\left(m\frac{n}{N}\right)\frac{\overline{E}}{P}$$

Equation (4) constitutes the baseline specification for our empirical analysis. Comparative statics exercises based on (2) show that k, and hence cash withdrawals, is increasing in ATM diffusion (d^{ATM}) and decreasing in the diffusion of non-cash payments (d^{Q}).[9] This occurs as a greater ATM diffusion increases the probability of a successful ATM withdrawal, reducing its expected cost. Similarly, a greater diffusion of alternative payment networks (e.g. debit cards and POS) reduces their cost, thus lowering the amount of cash used for transactions. The effect of an increase in the nominal interest rate R is ambiguous.[10] In addition, (4) shows that cash withdrawals increase with the number of agents holding an ATM card (n).

More wealth, as proxied by the level of real consumption expenditure, raises the cost of time and hence the value of f, which has an ambiguous effect on the demand for cash. A sufficient condition for cash withdrawals to be positively related to the level of wealth is that the probability of a successful ATM withdrawal is sufficiently greater than that of non-cash payment. Under this assumption, the elasticity of cash withdrawals to expenditure (both expressed in real terms) is greater than one. As we shall see, this non-homogeneity is key in the empirical tests developed in Section 3.2.

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8 In Section 3.2 we actually estimate a nonlinear specification, $q = m\left(\frac{n}{N}\right)^b$, and fail to reject the null hypothesis b=1.

9 These predictions are implied by (2). Let $F(k, d^{ATM}, d^{Q}, R) = (R - f)(z + kz_2) + f(s + (1-k)s_2)$, where superscripts have been overlooked. Since F_k>0, as from the second order conditions, the implicit theorem function yields ($z_{1,2}$=$s_{1,2}$=0 is assumed): $\frac{\partial k}{\partial d^{ATM}} = \frac{z_1(f - R)}{F_k} > 0$; $\frac{\partial k}{\partial d^{Q}} = \frac{-s_1 f}{F_k} < 0$; $\frac{\partial k}{\partial R} = \frac{-(z + kz_2)}{F_k}$ (sign is ambiguous), $\frac{\partial k}{\partial f} = \frac{(z + kz_2) - (s + (1-k)s_2)}{F_k}$ (sign is ambiguous).

10 For a given value of z, an increase in R reduces the demand for ATM cash since it increases its cost. When the proportion of ATM cash payments (k) is sufficiently high, however, this effect is offset (and can even change sign) due to the fact that increasing k reduces the probability of a successful ATM withdrawal, which raises the expected cost of cash.

3. Empirical evidence

3.1 A preliminary look at the data

Our dataset comprises quarterly time series over the 1993.II-2003.IV period.[11] Figure 2 shows that during the first three quarters of 2002 the average unit cash withdrawal from ATM records a sharp increase (from 157 to 166 euro), but only after an equally sharp fall in 2001 (from 165 to 157 euro). Overall, the unit withdrawal after the changeover remains around the same values recorded in the previous five years.[12] The same conclusion holds for the frequency of ATM usage (solid line), roughly constant at 6.3 withdrawals per card per quarter since 1995. Figure 2 also displays the behavior of households' real consumption of non durable goods and services in the period. Growth begins to slow down in the second half of 2000, bottoming out in the fourth quarter of 2001. Overall, annual consumption in 2002 remains unchanged as compared to 2001 (in 2001 it increased by just 0.1 per cent relative to 2000).

At least *prima facie*, none of these time series displays a behavior that might signal a sharp increase in the price level after the introduction of the euro banknotes and coins. However, this descriptive evidence can be potentially misleading, for at least two reasons. The first relates to our measure of consumption. Assume that inflation perceptions are correct, and that the official nondurable goods deflator underestimates inflation in 2002-2003; then, real consumption growth in Figure 2 should be correspondingly overestimated. In other words, real consumption could have slumped in reaction to a price hike. Second, the descriptive evidence does not take into account several important structural changes occurred in the last 15 years: first and foremost, the diffusion of ATM terminals over the national territory, and the related increase in aggregate cash withdrawals, but also other factors, such as the spread of POS terminals and the related increase in the use of debit cards. These developments typically follow long-term trends, and therefore are unlikely to have obscured the effect of the hypothesized price jump on the demand for cash in 2002-03. However, the model of Section 2 suggests that they can in principle have an important effect on the demand for ATM withdrawals, and should be taken into account in a rigorous analysis. The next subsection attempts to address these problems by estimating an equation that links the aggregate value of ATM withdrawals to its determinants.

[11] We refer to two sources. The flow of cash withdrawn from ATMs in Italy, the number of ATM cards, POS and ATM terminals, and the interest rate on checking accounts are provided to the Bank of Italy by the banking system for supervisory reasons. Data on consumption of services and non-durable goods, and the related deflators, are released by the Italian national statistical office (Istat). All the series used in the paper refer to Italy.

[12] Such stationarity, in a period of moderate but positive inflation, might reflect the increasing use of cash substitutes. The regression analysis below supports this hypothesis.

3.2 Econometric analysis: inference from an estimated currency expenditure equation

Taking logs of both sides, equation (4) becomes:

$$(5) \qquad c_t - p_t = g_0 + g_1 d_t^{ATM} + g_2 d_t^{Q} + g_3 (e_t - p_t) + g_4 r_t + g_5 n_t + e_t$$

where lowercase letters denote logs, and e_t is an error term with variance s_e^2. We measure c_t by the nominal value of nationwide quarterly withdrawals from ATM terminals, n_t by the number of outstanding ATM cards (at the end of the quarter). The diffusion of the ATM network, d^{ATM}, is proxied by the ratio between the number of ATM terminals nationwide and n_t. Similarly, the diffusion of alternative payment instruments, d^{Q}, is measured by the ratio between the number of POS terminals and n_t. Finally, e_t and p_t are proxied by the aggregate nominal consumption of services and non durable goods, and by its deflator, in the order, whereas r_t is the interest rate on checking accounts.[13]

Can equation (5) help us shed light on the issue at the core of this paper? In what follows we spell out the assumptions and the econometric requirements needed for the answer to be positive. One key assumption is that until the fourth quarter of 2001 all of the time series appearing in (5) are measured with no error. In particular, p_t, the true (log) price level, coincides with p_t^o, the observed deflator, measured by the national statistical office. After the changeover, we allow for the possibility that the official deflator may be affected by measurement error, and continue to assume that all other variables are correct. Under this assumption (which appears reasonable, since until the end of 2001 there was no argument about data quality) the coefficient of (5), if estimated over the pre-changeover period, will not be affected by measurement problems.

We therefore begin by estimating (5) using data until 2001.IV. The results are reported in Table 1, column *(a)*.[14] The estimate is in line with the suggestions of the theory: the diffusion of non-cash forms of payment reduces cash withdrawals, while the diffusion of ATM terminals increases them, as does consumption expenditure. The coefficient of the interest rate on checking accounts is negative but not significant.

[13] Based on (4) the population size, *N*, should also appear among the regressors. We omit it because it was roughly constant over the estimation period.

[14] Three quarterly dummies (not shown) are included among the explanatory variables to account for seasonal effects.

Table 1: Estimates of equation (5)

	estimation period				
	1993.II-2001.IV	*1993.II-2003.IV*			
		test of (6)		*test of (6')*	
	(a)	*(b)*	*(c)*	*(d)*	*(e)*
ATM terminals diffusion (d_t^{ATM})	2.38**	2.38	2.34**	2.38	2.43**
	5.9	Constr.	7.0	Constr.	6.8
POS terminals diffusion (d_t^Q)	-17.05**	-17.05	-17.00**	-17.05	-15.75*
	-5.0	Constr.	-6.9	Constr.	-5.7
Real consumption (log; e_t- p_t)	2.82**	2.82	2.75**	2.82	2.57**
	3.6	Constr.	4.0	Constr.	3.8
Number of ATM cards (log; n_t)	0.89**	0.89	0.91**	0.89	0.89**
	5.6	Constr.	6.5	Constr.	6.2
Interest rate on deposits (log; r_t)	-7.1e-4	-7.1e-4	3.4e-3	-7.1e-4	6.9e-3
	0.0	Constr.	0.2	Constr.	0.3
Constant	-8.82	-8.82	-8.57	-8.82	-5.42
	-0.8	Constr.	-1.1	Constr.	-0.6
Dummy 2002.I-03.IV	-	-6.2e-3	-6.4e-3	-	-
	-	-0.6	-0.4	-	-
Dummy 2002.I-03.IV*linear trend	-	-	-	-2.0e-3	-3.6e-3
	-	-	-	-1.0	-1.1
s_e^2	-	8.8e-04	8.9e-04	8.8e-04	8.9e-04
$s_h^2 = s_e^2 + (1-g_3)^2 s_x^2$	-	3.5e-04	2.5e-04	2.7e-04	1.5e-04
F test for $s_e^2 \leq s_h^2$	-	2.52	3.61*	3.24	5.89**
N° observations	35	43	43	43	43
R^2	0.98	-	0.99	-	0.99
DW	1.47	1.59	1.58	1.60	1.60

Note: The dependent variable is the (log of) aggregate cash withdrawals from ATM in real terms. s_e^2 and s_h^2 denote, in order, the variance of the error term in equation (5) before and after 2001.IV. OLS estimates. Heteroskedasticity-robust *t*-statistics are reported below each coefficient. One or two asterisks denote, respectively, 5 and 1 per cent significance levels. The regressions also include three seasonal dummies (coefficients not reported). The linear trend takes integer values between 1 and 8 over the 2002.I – 2003.IV period.

Altogether, while simple, equation (5) seems to capture some essential features of the demand for ATM withdrawals. Considering that it does not feature a lagged dependent variable on the right-hand side (we experimented with specifications incorporating one, but the related coefficient turned out to be non significant), it tracks the data quite well.

We can now test the null hypothesis that after the changeover $p_t = p_t^o$, against the alternative $p_t > p_t^o$. Since p_t is not observable after 2001.IV, we consider two hypotheses about its behavior. The first is the following:

$$(6) \qquad p_t = p_t^o, \; t \leq 2001.\text{IV}, \qquad p_t = p_t^o + a + x_t, \; t \geq 2002.\text{I}$$

where a is a positive constant and x_t a white noise term independent of e_t, with variance s_x^2. Expressions (6) could be an appropriate description of a one-off increase in the true price level after the changeover. A second hypothesis is:

$$(6') \qquad p_t = p_t^o, \; t \leq 2001.\text{IV}, \qquad p_t = p_t^o + gT + x_t, \; t \geq 2002.\text{I}$$

where g is a positive constant and T is a linear trend such that T=1 in 2002.I. This formulation would entail a widening gap between the observed (official) and the true price deflator, implying a permanent inflation bias. It would probably be unrealistic for large T, but could be appropriate over our sample period, which only covers eight quarters after the changeover.

Substituting (6) into (5) yields:

$$(7) \qquad c_t - p_t^o = g_0 + q_0 + g_1 d_t^{ATM} + g_2 d_t^Q + g_3(e_t - p_t^o) + g_4 r_t + g_5 n_t + h_t, \qquad t \geq 2002.\text{I},$$

where $q_0 = (1-\gamma_3)a$, and $h_t = e_t + (1-\gamma_3)x_t$, with variance $s_h^2 = s_e^2 + (1-g_3)^2 s_x^2$. A way to test the null hypothesis of no distortion in the price level after the changeover against the alternative hypothesis (6) would then entail estimating equation (7) over the entire sample period 1993.II-2003.IV, introducing a dummy variable to allow the constant to change over the last two years, and checking for heteroskedasticity. However, it is easy to check that under the alternative hypothesis (6), equation (7) is affected by a classic errors in variables problem; if $s_x^2 > 0$, OLS coefficients would be inconsistent and biased towards zero. To circumvent this difficulty, we restrict the g_i to take the values estimated over the 1993.II-2001.IV period, and use the 2002.I-2003.IV data to estimate the coefficient of the 2002-03 dummy, q_0, which is an unbiased estimator of $(1-\gamma_3)a$. The results of this exercise are presented in column *(b)* of Table 1. The estimated q_0 has a value of –0.0062, implying that the average inflation rate in 2002 was 0.3 percentage points higher than computed using the official deflator. However, a one-tail t test cannot reject the null that the coefficient is zero against the alternative that it is negative. Note that for the test to have power against the alternative it is essential that γ_3, the elasticity of real consumption, be different from 1, i.e. that equation (5) be non homogeneous of degree one in the deflator p_t, as predicted by the model in Section 2. Indeed, the estimated g_3 is 2.82, statistically different from 1 at the 5 per cent level ($F_{(1,26)}$=5.02, p-value of 0.03). Column *(b)* also reports the estimated s_e^2 and s_h^2. The hypothesis $s_h^2 \geq s_e^2$, which should hold based on (6), is not statistically rejected, but the former variance is smaller than the latter by a factor of 3.

As mentioned above, if (6) were true and the error-in-variables problem were serious after 2001.IV, OLS coefficients should be biased towards zero. Therefore, as a further check we estimate all the parameters of (7) over the entire sample period. The related results, in column *(c)*, show that the parameters remain remarkably stable. In this case the hypothesis $s_h^2 \geq s_e^2$ can be rejected at the 5 percent confidence level.

Next, we replicate the exercise for our second alternative hypothesis. Substituting (6') in (5) yields an equation identical to (7), except for the fact that now $q_0 = (1-g_3)gT$. Thus, beginning with 2002.I a linear trend with coefficient $g(1-g_3)$ should enter the equation. Also, the error term should display the same form of heteroskedasticity as under hypothesis (6). Specifications in columns *(d)-(e)* of Table 1 show, as before, no evidence consistent with the hypothesis of an increase in the price level after the changeover. In both cases the coefficient of the 2002.I-2003.IV dummy interacted with the time trend is negative but not statistically different from zero; in *(e)* the hypothesis $s_h^2 \geq s_e^2$ is strongly rejected.[15,16]

[15] Since the exercises just described entail detecting a structural break in the equation after the fourth quarter of 2001, it is important that the coefficients in column *(a)* of Table 1 be stable. An obvious candidate for a structural break is the beginning of the single euro area monetary policy regime, in January 1999. Therefore, the five coefficients $\hat{g}_0$ through $\hat{g}_5$ of specification *(a)* were allowed to change over the 1999.I-2001.IV period. The F test of the null hypothesis that they are jointly equal to zero yields $F_{(5,20)}$ =2.66, which does not allow to reject the null of parameter stability at the 5 percent level. However, since this value is close to significance, the tests were replicated with the equation incorporating these extra terms. The results were qualitatively analogous to those in Table 1 and in figure 3, and therefore are not reported.

[16] While we focused on hypotheses (6) and (6') in order to maximize the power of the test, we also tested hypothesis $p_t = p_t^o + a + gT + x_t$, $t \geq$2002.I, which nests (6) and (6'). The estimated coefficients for the 2002-03 dummy and for the time trend results were not statistically different from zero (individually as well as jointly). It is also worth mentioning that the results in this section are robust to changes in the specification of equation (5). The specification used in Angelini, Ardizzi and Lippi (2005) features inflation and a time trend among the regressors, whereas expenditure is proxied by consumption of non durable goods.

3.3 Exploring the power of the statistical test

What is the power of the tests performed above? It could be argued that the equations in Table 1 fail to detect a price increase due to lack of power. To investigate this hypothesis, we perform a counterfactual exercise. We assume that beginning in 2002.I the true price deflator is higher than the official deflator. Using (7), we set $p_t^o = p_t - a$ after the changeover. We then assign numeric values to a and re-estimate specifications *(b)* and *(c)*. If our tests have sufficient power, the coefficient of the 2002-03 dummy should become negative and significant for relatively small values of a. We report the results of this exercise for values of a ranging between 0.005 and 0.1, implying that in 2002 inflation was between 0.5 and 10 percentage points higher than recorded by official statistics.

Figure 3 plots the "true" a, measured on the horizontal axis, against its estimated value $\hat{a}$, obtained as the ratio between the coefficient of the 2002-03 dummy and (1-$\hat{g}_3$). The curves obtained with specifications *(b)* and *(c)* virtually overlap, so that they can hardly be distinguished in the figure. They are very close to the 45° lines, indicating that the size of the distortion is captured quite well - in fact, it is systematically slightly overestimated. The figure also shows the precision of the estimates, measured by the t statistic of the 2002.I-2003.IV dummy. Specification *(b)* manages to correctly signal as statistically significant (with a 5 percent confidence level) a value of a as little as 1 percent; in this case specification *(c)* yields a p-value of .09. Both specifications capture values of a greater than 1 percent at least at the 5 percent level.

4. Conclusions

Did the euro cash changeover trigger a sudden, substantial increase of the price level in the euro area, largely undetected by the national statistical offices? This paper presented a simple indirect method to address this question. The basic idea underlying the testing strategy entails searching for the effects that the hypothesized large increase in the price level should have induced on the dynamics of payment instruments, notably cash withdrawals from the ATM network.

The simple theoretical model developed in the paper highlights some determinants of ATM withdrawals (ATM and POS network diffusion, expenditure and real wealth) and suggests that the relationship between withdrawals and expenditure is not homogenous of degree one in the price level. The latter prediction is key in the empirical analysis because it allows implications on the price level to be derived from the dynamics of the observed time series.

The estimation exercise, conducted along the lines suggested by the theory, confirms that the cash demand equation is indeed strongly non homogeneous with respect to the price level. This allows us to set up econometric tests of the null hypothesis that after the currency changeover, in the first months of 2002, the official price index continued to correctly measure actual price dynamics, against the alternative that it underestimated it. The main result of the analysis is that none of the various tests performed provides evidence against the null.

To assess the possibility that failure to reject the null is due to lack of power, we perform a counterfactual exercise: we introduce an artificial increase in the deflator time series beginning in 2002, and re-estimate the demand equation over the 2002-03 period, allowing for parameter instability. The equation accurately captures the magnitude of the inflation distortion, correctly signaling it as statistically significant as soon as it grows greater than or equal to 1 percent on an annual basis. We conclude that the determinants of the well-documented disconnect between inflation as perceived by consumers and as measured by the national statistical offices of the euro area countries cannot be ascribed to a mismeasurement by the latter.

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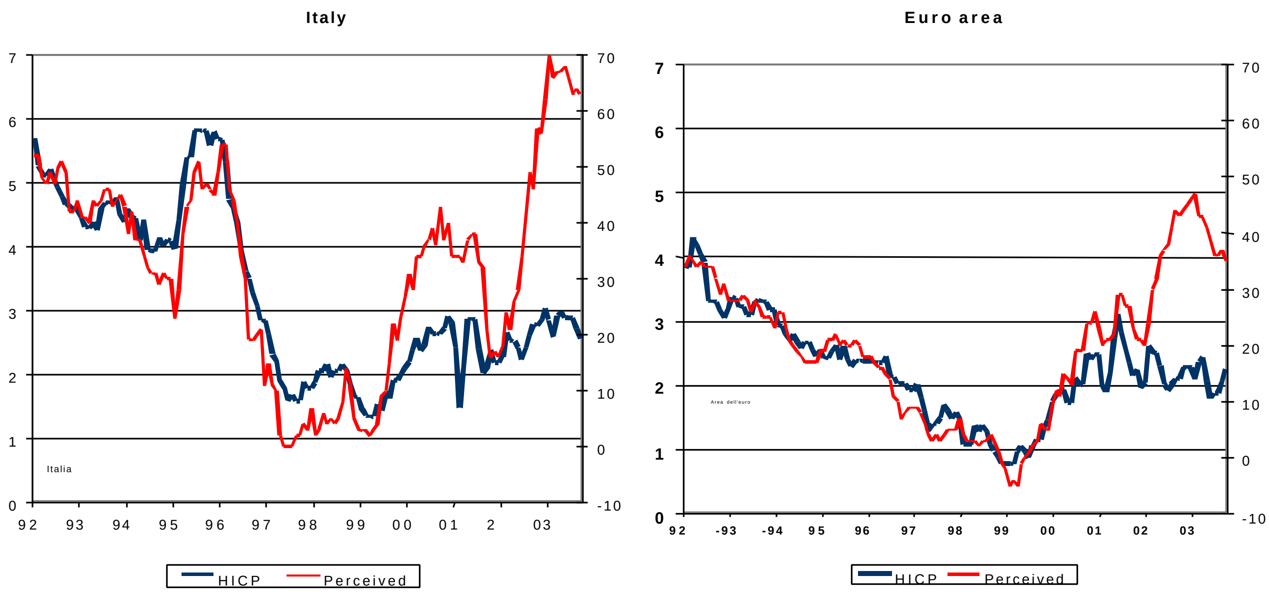
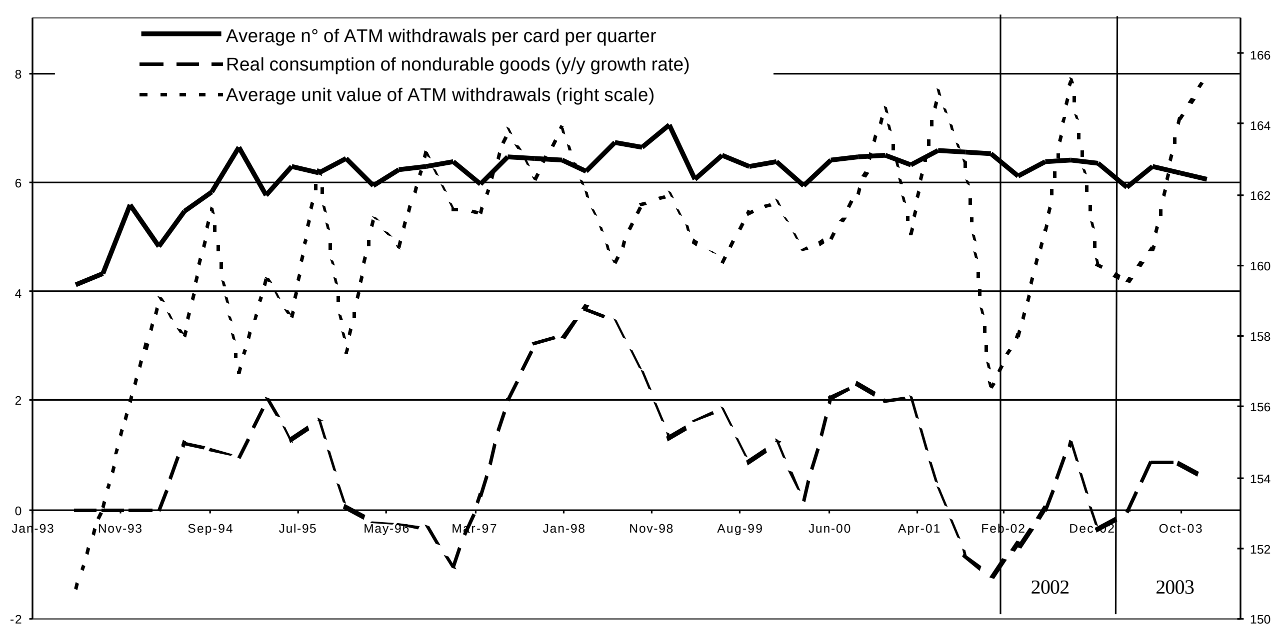
Fig. 1: Perceived and official inflation in selected euro area countries



Source: ISTAT; European Commission.

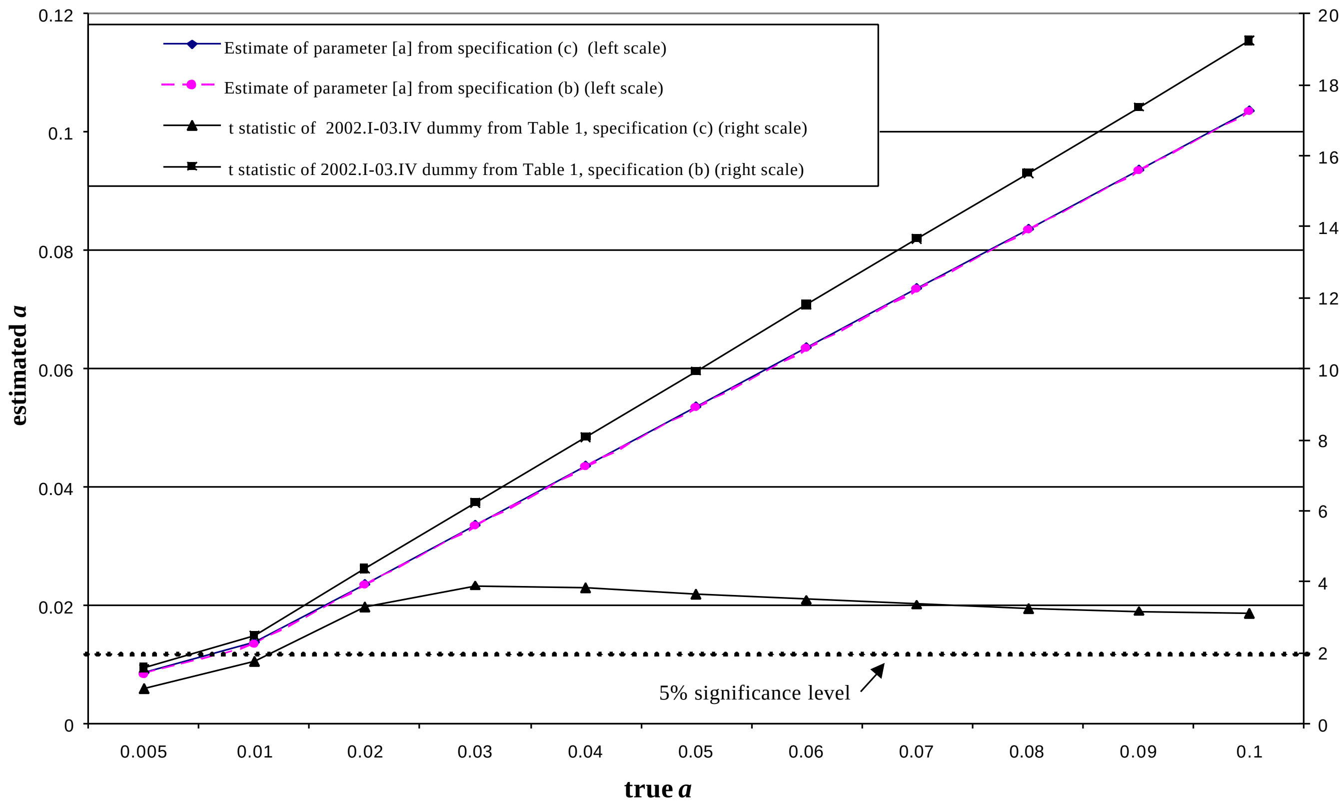
Note: HICP stands for inflation measured by twelve-month growth rate of the Harmonized Index of Consumer Prices (HICP; left axis). Perceived stands for perceived inflation based on surveys by ISAE and by the European Commission. It is computed as the difference between the share of respondents reporting that prices "strongly increased" or "moderately increased" and the share of respondents reporting "stable" or "decreased" prices (right axis).

Fig. 2: ATM usage and consumption growth (quarterly data)



Source: Bank of Italy; ISTAT.

Fig. 3: The power of the statistical test: Results from a counterfactual exercise



Note: Equations *(b)* and *(c)* in table 1 were re-estimated using a counterfactual price deflator, which jumps by a in 2002.I. See the text for a detailed description of the exercise. The horizontal axis measures a, the shift in the price level occurred after the changeover, based on hypothesis (6) in the text. The vertical axis measures the estimated a, computed as: $\hat{a}$ = (estimated coefficient of 2002.I-2003.IV dummy)/(1-$\hat{g}_3$). The t statistics are reported in absolute value.