BANCA D'ITALIA

Temi di discussione

del Servizio Studi

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Indirect evidence from ATM withdrawals**

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Number 581 - March 2006

DID INFLATION REALLY SOAR AFTER THE EURO CASH CHANGEOVER? INDIRECT EVIDENCE FROM ATM WITHDRAWALS

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Abstract

The introduction of the euro notes and coins in the first two 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 behavior 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 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: E41, E31, E51.

Keywords: banknotes, currency, euro, inflation.

Contents

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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. While the quotations reported above refer to 2002, the perception that the euro brought about higher inflation is still vivid at present. [1]

It is somewhat puzzling that a change in the unit of account might have an impact on the inflation rate. Indeed, a number of conjectures have been formulated to explain the discrepancy between inflation perceptions and the official statistics, emphasizing the role of psychological factors (e.g. Traut-Mattausch *et al.*, 2004) and/or the disproportionate influence of a few industry prices on individual perceptions. Hobijn, Ravenna and

[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. Two years later, 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 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).

Tambalotti (2004) and Gaiotti and Lippi (2004) analyze the dynamics of restaurant prices and find evidence consistent with a price hike (mainly driven by a lumping of price revisions in an industry where price revisions are normally infrequent). Deutsche Bundesbank (2004) provides comparable evidence for some German services (restaurants, cinemas, dry-cleaning and hairdressers). Other papers argue that inflation perceptions are mainly affected by the prices of goods that are cheaper and more frequently purchased (Del Giovane and Sabbatini, 2005; Ehrmann, 2005). Dziuda and Mastrobuoni (2005) and Mastrobuoni (2004) present a model that rationalizes why such goods are the ones that actually record greater price increases. Still, these studies do not provide a direct answer on whether the *general* price index was measured with error during the changeover. Rather, they maintain the assumption that official statistics are correct. The main obstacle faced by researchers interested in verifying this assumption is the absence of reliable alternative inflation measures. The thesis that price increases were much larger than measured by the national statistical offices, suggested by the indicators of perceived inflation, remains mostly based on anecdotal evidence.

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), at the beginning of January 2002, by using data on currency withdrawn from the ATM network. 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 consumer expenditure is not homogenous of degree one 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 ATM withdrawals and consumer 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 sample to the changeover period (2002-03) should cause a specific form of 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. We show that our test is sufficiently powerful to identify a bias greater than 0.5 percentage point. The analysis fails to find evidence consistent with the occurrence of a price hike after the changeover.

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 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: in the twelve months ending in December 2001 currency in the euro area experienced an unprecedented contraction (-32,5 percent; see e.g. European Central Bank, 2002b; a similar figure was recorded in Italy). 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 deeply influenced by portfolio reasons, as repeatedly argued by the European Central Bank. By contrast, there is no obvious reason why ATM withdrawals – undoubtedly mainly driven by transactions demand – should have been affected by these same factors. The raw data in Figure 2 broadly confirm this a priori: neither the average number nor the unit value of ATM withdrawals made in each quarter by a typical cardholder show the dramatic signs of discontinuity which are clearly evident in the time series of the narrow monetary aggregates.

The paper is organized as follows. In the next Section we present a model of the 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 cash withdrawals

This section presents a simple model of cash withdrawals that provides guidance for the empirical analysis of Section 3, where data on ATM withdrawals aggregated at the national level are used. To match these data, we first focus on the choice between cash and alternative means of payments by a representative ATM cardholder. Next, we present an aggregation to account for the growing number of cardholders over our estimation period.[2]. As we shall see, this second step introduces a few explanatory variables in our cash withdrawal equation which are typically not present in a money demand function.

Let i index an agent who possesses an ATM card and 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, etc.[3] 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.

The cost of cash substitutes is normalized to zero and R is used to denote the cost of cash relative to "substitutes" (per unit of expenditure). This can be thought of as the nominal interest rate, whose value determines the amount of forgone interest on deposits. Using a substitute payment instrument, e.g. 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 ($1–z$) the agent fails to withdraw and bears a cost ϕ, $\phi>R$, 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

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

[3] 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.

wealthier people, ϕ is assumed to be an increasing function of individual wealth (proxied by real consumption expenditure). Thus $\phi\left(E^i / P\right)$ has $\phi_1 > 0$ (P denotes the price level). 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 ϕ.[4]

When deciding whether to use cash or "substitutes" to carry out transactions, the agent solves the following problem:

$$\text{(1)} \qquad \underset{k}{\text{Min}} \quad k\,[R\,z(d^{ATM},k)+(1-z(d^{ATM},k))\phi] \;+\; (1-k)\,[1-s(d^Q,1-k)]\phi$$

where i superscripts have been overlooked and 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:[5]

$$\text{(2)} \qquad (R-\phi)(z+kz_2)+\phi(s+(1-k)s_2)=0$$

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

$$C^i = k(d^{ATM}, d^Q, \phi, R)\, E^i$$

A comparative statics exercise based on (2) shows that k is increasing in ATM diffusion (d^{ATM}) and decreasing in the diffusion of non-cash payments (d^Q). 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

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[4] To keep matters simple we assume that the agent incurs the cost ϕ 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*).

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

cards and POS) reduces their cost, thus lowering the amount of cash used for transactions. An increase in the nominal interest rate *R* decreases the demand for cash since it increases its opportunity cost. Finally, the model implies that the proportion of cash expenditures is greater if the opportunity cost of time increases (ϕ): intuitively, wealthier people hold relatively more cash since the cost of incurring into a malfunctioning ATM is greater for them.[6]

Let us now consider the aggregation problem. Let n be the number of agents with an ATM card and N be the country population size. Denote aggregate withdrawals and expenditure for the group of ATM cardholders by, respectively, $C=nC^i$ and $E=nE^i$. Further, let $\theta \equiv E/\overline{E}$, where $\overline{E}$ denotes aggregate country-wide expenditure. In order to bring equation (3) to the data, we model θ as $\theta = \mu \frac{n}{N}$.[7] Thus, $E^i = \frac{\mu}{N}\overline{E}$. 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:

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

Equation (4) naturally shows that aggregate withdrawals increase with the number of agents holding an ATM card (n). More interestingly, it implies that the elasticity of cash

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[6] These predictions are obtained from (2). At an internal solution the following condition holds: $1 > \frac{\phi - R}{\phi} = \frac{(s+(1-k)s_2)}{(z+kz_2)} > 0$, which implies: $(z+kz_2) > 0$, $(s+(1-k)s_2) > 0$ and $(z+kz_2) > (s+(1-k)s_2)$. Omit the i-superscript and let $\mathrm{F}(k, d^{ATM}, d^{Q}, R, \phi) = (R-\phi)(z+kz_2) + \phi(s+(1-k)s_2)$. Since $F_k>0$ (from the second order conditions), the implicit function theorem yields (assume $z_{1,2}=s_{1,2}=0$): $\frac{\partial k}{\partial d^{ATM}} = \frac{z_1(\phi - R)}{F_k} > 0$; $\frac{\partial k}{\partial d^{Q}} = \frac{-s_1\phi}{F_k} < 0$; $\frac{\partial k}{\partial R} = \frac{-(z+kz_2)}{F_k} < 0$, $\frac{\partial k}{\partial \phi} = \frac{(z+kz_2)-(s+(1-k)s_2)}{F_k} > 0$.

[7] In Section 3.2 we actually estimate a nonlinear specification, $\theta = \mu\left(\frac{n}{N}\right)^{\beta}$, and fail to reject the null hypothesis β=1.

withdrawals with respect to expenditure is greater than one,[8] due to the fact that the agents' proportion of cash expenditures is increasing in the opportunity cost of time (ϕ). As we shall see in Section 3.2, this prediction is clearly borne out by the data.

3. Empirical evidence

This section begins by presenting some descriptive evidence on ATM withdrawals during the period 1993-2003. We then use the model of Section 2 to estimate a currency expenditure equation and formally test for the presence of measurement error in the aggregate price level. We conclude the section exploring the power of our statistical test and analyzing the robustness of our results (considering e.g. parameter instability and potential endogeneity of the regressors).

3.1 A preliminary look at the data

Our dataset comprises quarterly time series over the 1993.II-2003.IV period.[9] 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 (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). 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

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[8] Recall that $\frac{\partial k}{\partial \phi} > 0$ and $\phi_1 > 0$ (i.e. the cost of time is increasing in wealth). Taking logs of (4) it follows that the elasticity is: $\varepsilon = \frac{\partial \log(C/P)}{\partial \log(\overline{E}/P)} = 1 + \frac{\partial k}{\partial \phi}\frac{\phi_1}{k}\frac{\mu}{N}\frac{\overline{E}}{P} > 1$.

[9] 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.

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.

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.

3.2 Econometric analysis: inference from an estimated currency expenditure equation

A log-linear version of equation (4) is:

$$(5) \qquad c_t - p_t = \gamma_0 + \gamma_1 d_t^{ATM} + \gamma_2 d_t^{Q} + \gamma_3 (e_t - p_t) + \gamma_4 r_t + \gamma_5 n_t + \varepsilon_t$$

where lowercase letters denote logs and ε_t is an error term with variance σ_ε^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 the cost of cash relative to substitutes is proxied by the interest rate on checking accounts (r_t).[10]

Equation (5) shows that as long as the elasticity of cash withdrawals with respect to real consumption (γ_3) is different from one then price level dynamics can be deduced from the equation based on ATM withdrawals and consumption data. Can equation (5) then 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 measured correctly. 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)*.[11] The coefficients are 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. The elasticity of cash demand to consumption expenditure, γ_3, is 2.82, statistically greater than one at the 5 percent level (t statistic of 2.2). The coefficient of the interest rate on checking accounts is negative but not significant. The coefficient on the number of ATM cards is .88, not significantly different from one (t statistic of 0.7).

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

[10] 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.

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

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) $p_t = p_t^o$, $t \leq 2001.\text{IV}$, $p_t = p_t^o + a + \xi_t$, $t \geq 2002.\text{I}$

where a is a positive constant and ξ_t a white noise term independent of ε_t, with variance σ_ξ^2. Expressions (6) could be an appropriate description of a one-off increase in the true price level after the changeover. An alternative hypothesis is:

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

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) $c_t - p_t^o = \gamma_0 + \theta_0 + \gamma_1 d_t^{ATM} + \gamma_2 d_t^Q + \gamma_3 (e_t - p_t^o) + \gamma_4 r_t + \gamma_5 n_t + \eta_t$, $t \geq 2002.\text{I}$,

where $\theta_0 = (1-\gamma_3)a$, and $\eta_t = \varepsilon_t + (1-\gamma_3)\xi_t$, with variance $\sigma_\eta^2 = \sigma_\varepsilon^2 + (1-\gamma_3)^2 \sigma_\xi^2$. 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.

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
σ_ε^2	-	8.8e-04	8.9e-04	8.8e-04	8.9e-04
$\sigma_\eta^2 = \sigma_\varepsilon^2 + (1-\gamma_3)^2 \sigma_\xi^2$	-	3.5e-04	2.5e-04	2.7e-04	1.5e-04
F test for $\sigma_\varepsilon^2 \le \sigma_\eta^2$	-	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. σ_ε^2 and σ_η^2 denote, in order, the variance of the error term in equation (7) 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.

However, it is easy to check that under the alternative hypothesis (6), equation (7) is affected by a classic errors in variables problem; if $\sigma_{\xi}^{2}>0$, OLS coefficients would be inconsistent and biased towards zero. To circumvent this difficulty, we restrict the γ_i to take the values estimated over the 1993.II-2001.IV period, and estimate only the coefficient of the 2002-03 dummy, θ_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 θ_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 one, i.e. that equation (5) be non homogeneous of degree one in the deflator p_t, as predicted by the model in Section 2 and confirmed by the estimates in column *(a)*. Column *(b)* also reports the estimated σ_{ε}^{2} and σ_{η}^{2}. The hypothesis $\sigma_{\eta}^{2}\geq\sigma_{\varepsilon}^{2}$, which should hold based on (6), is not statistically rejected (note however that σ_{η}^{2} is smaller than σ_{ε}^{2}).

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 $\sigma_{\eta}^{2}\geq\sigma_{\varepsilon}^{2}$ 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 $\theta_0 = (1-\gamma_3)gT$. Thus, beginning with 2002.I a linear trend with coefficient $g(1-\gamma_3)$ 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 $\sigma_{\eta}^{2} \geq \sigma_{\varepsilon}^{2}$ is rejected.[12]

3.3 Exploring the power of the statistical test

Our econometric procedure amounts to a *t* test of the coefficient of a dummy in a linear regression. Therefore, the properties of our tests and their statistical power are well grounded in standard asymptotic and small sample theory. It could be argued, however, that the precision of our estimates is insufficient to generate adequate power, e.g. because of the short sample period available. 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 one. Using (6), 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-$\hat{\gamma}_3$). 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. Both specifications fail to detect the presence of a 0.5 percentage point distortion. However, specification *(b)* manages to correctly signal as statistically significant (with a 5 percent confidence level) a value of a as little as one percent; in this case

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[12] While we focused on hypotheses (6) and (6') in order to maximize the power of the test, we also tested hypothesis $p_t = p_t^o + a + gT + \xi_t$, $t \geq 2002.I$, 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).

specification *(c)* yields a p-value of .09. Both specifications capture values of *a* greater than one percent at least at the 5 percent level.

3.4 Robustness check

The above results were subjected to a number of robustness checks. First, we checked the stability of the specification reported in Table 1, column *(a)*. The exercises described in section 3.2 entail detecting a structural break in the equation after the fourth quarter of 2001. Thus, 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 $\hat{\gamma}_0$ through $\hat{\gamma}_5$ of specification *(a)* were allowed to change over the 1999.I-2001.IV period. The F test of the null hypothesis that the changes in the coefficients 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 in Table 1 were replicated with the equation incorporating these extra terms. The results were qualitatively analogous to those in Table 1, and therefore are not reported.

Second, the estimates of equation (5) reported in Table 1 are subject to a potential endogeneity bias problem, as some right hand-side variables (e.g. expenditure) may be simultaneously determined with the dependent variable. Thus, we re-estimated specification *(a)* with two-stage least squares, instrumenting d_t^{ATM}, d_t^{Q}, e_t-p_t and n_t with their lagged values. The results (not reported) are virtually unchanged.

Third, the analysis in section 3.2 relies on the hypothesis that nominal expenditure e_t is measured without error prior and after the changeover. However, nominal components of consumption expenditure are computed using both value data (i.e. data measured in nominal terms) and data built from price and quantity indices. Possible mis-measurements in the prices of these components after 2001.IV will in principle bias e_t as well. Since a detailed breakdown of the data on household consumption by construction method is not available, it is not possible to address this concern in a precise way (see ISTAT, 2000). However, ISTAT does publish a breakdown of consumption expenditure used in section 3.2 into two categories: non durable goods and services. It turns out that the nondurables component is virtually entirely built from value data, and therefore it is not affected by possible mis-measurement in p_t after the changeover. Thus, we re-run the regressions in Table 1 proxying e_t with consumption of nondurables, i.e. excluding expenditure on services.[13] Table A1 in the appendix reports the instrumental variables estimates of this last specification. The elasticity of money demand to real consumption is now 2.30, slightly lower than in Table 1, but still significantly different from one at the 5 percent level. No appreciable changes in the other results emerge. Similar results were obtained with OLS.

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 and tested using a simple cash demand equation. 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

[13] Angelini, Ardizzi and Lippi (2005) also use consumption of nondurables as a proxy for expenditure, in a specification featuring inflation and a time trend among the regressors. The results of the tests are analogous to those reported here.

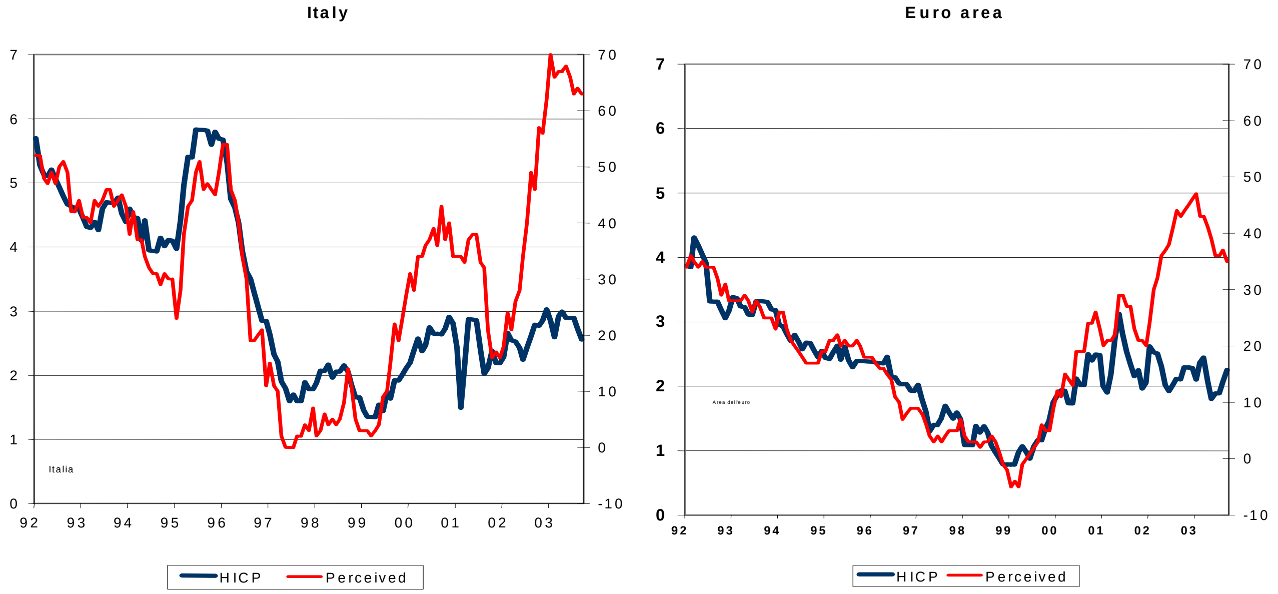
main result of the analysis is that none of the various tests performed provides evidence against the null.

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-run our tests. 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 one 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 mis-measurement by the latter.

Tables and figures

FIG. 1: PERCEIVED AND OFFICIAL INFLATION IN SELECTED EURO AREA COUNTRIES

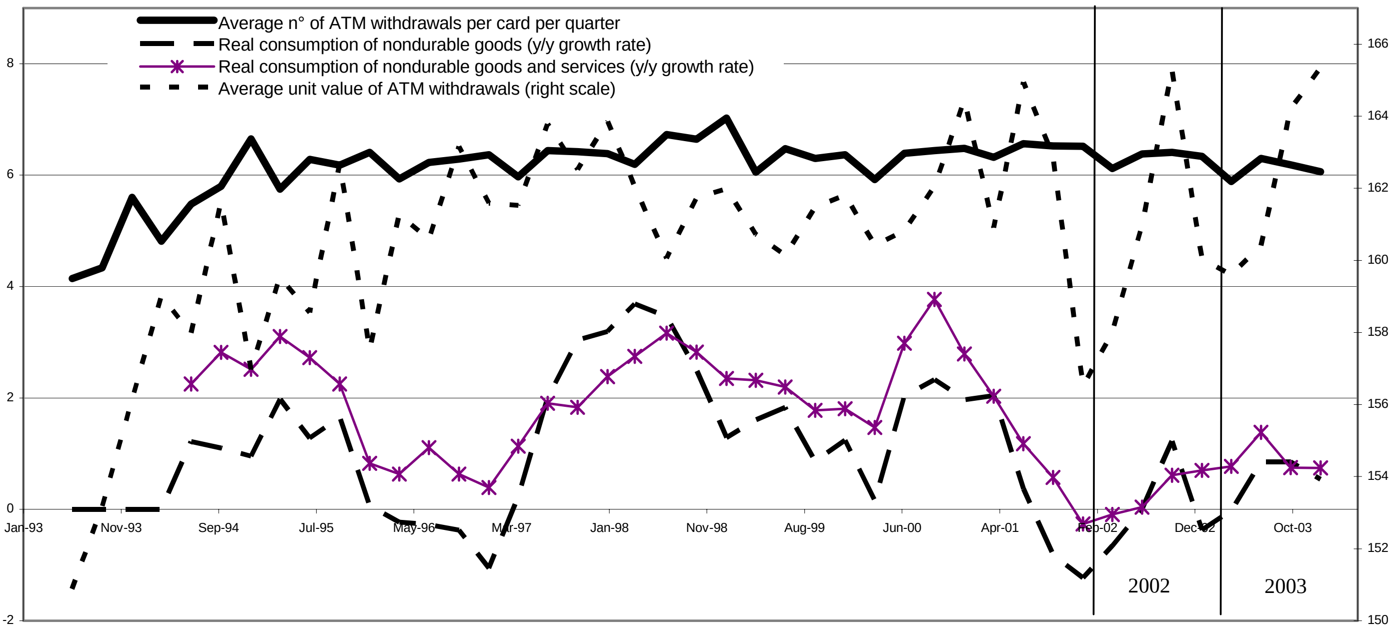
(monthly data)



Source: Eurostat, European Commission.

Note: HICP stands for inflation measured by twelve-month growth rate of the Harmonized Index of Consumer Prices (HICP; left axis). “Perceived” in the legend (right axis) stands for perceived inflation based on surveys by ISAE and by the European Commission, as reported in the monthly press release “Business and consumer survey results”. It is computed as the difference between the share of respondents reporting that prices “strongly increased” (weight 1) or “moderately increased” (weight 0.5) and the share of respondents reporting “stable” (weight 0.5)or “decreased” (weight 1) prices.

FIG. 2: ATM USAGE AND CONSUMPTION GROWTH

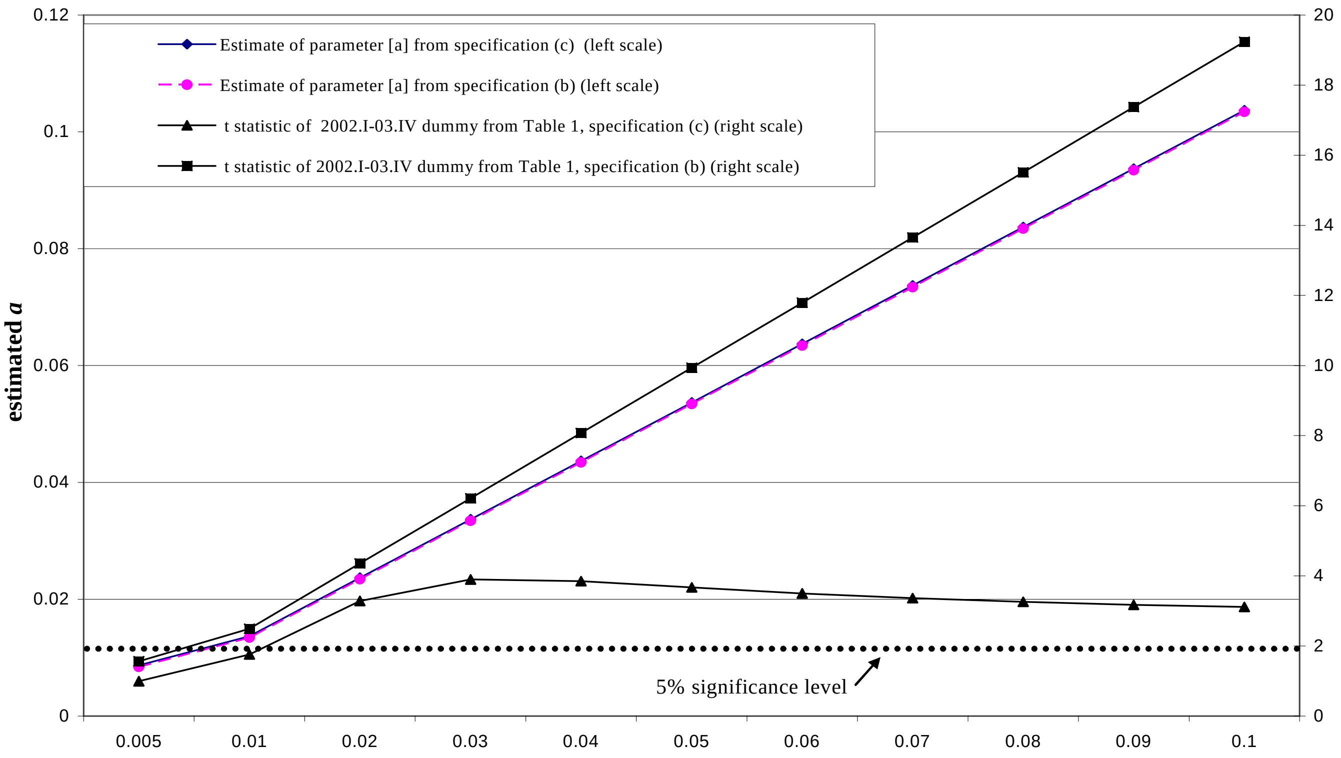


(quarterly data)

Source: Bank of Italy; ISTAT.

true

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-$\hat{\gamma}_3$). The t statistics are reported in absolute value.

Appendix

TABLE A1: ALTERNATIVE ESTIMATES OF EQUATION (5)

(e_t - p_t measured as consumption of nondurable goods; I.V. estimates)

	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})	2.47**	2.47	2.49**	2.47	2.54**
	4.1	Constr.	6.6	Constr.	4.9
POS terminals diffusion (d_t^{Q})	-12.49**	-12.49	-12.96**	-12.49	-12.13**
	-5.2	Constr.	-6.6	Constr.	-5.4
Real consumption (log; e_t - p_t)	2.30**	2.30	2.52**	2.30	2.30**
	3.9	Constr.	4.3	Constr.	3.7
Number of ATM cards (log; n_t)	1.11**	1.11	1.10**	1.11	1.08**
	7.7	Constr.	8.7	Constr.	8.7
Interest rate on deposits (log; r_t)	0.01	0.01	0.02	0.01	7.2e-3
	0.4	Constr.	0.8	Constr.	0.3
Constant	-4.40	-4.40	-6.56	-4.40	-3.54
	-0.6	Constr.	-1.1	Constr.	-0.5
Dummy 2002.I-03.IV	-	-8.5e-3	-3.7e-3	-	-
	-	-0.9	-0.3	-	-
Dummy 2002.I-03.IV*linear trend	-	-	-	-1.8e-3	-2.2e-3
	-	-	-	-1.0	-0.7
σ_ε^2	-	8.0e-4	7.8e-4	8.0e-4	7.8e-4
$\sigma_\eta^2 = \sigma_\varepsilon^2 + (1-\gamma_3)^2 \sigma_\xi^2$	-	3.3e-4	2.8e-4	3.1e-4	2.2e-4
F test for $\sigma_\varepsilon^2 \leq \sigma_\eta^2$	-	2.44	2.77	2.56	3.54*
N° observations	34	43	42	43	42
R^2	0.99	-	0.99	-	0.99
DW	1.85	1.69	1.88	1.69	1.88

Note: The dependent variable is the (log of) aggregate cash withdrawals from ATM in real terms. σ_ε^2 and σ_η^2 denote, in order, the variance of the error term in equation (7) before and after 2001.IV. Instrumental variables estimates. d_t^{ATM}, d_t^{Q}, n_t and e_t - p_t are instrumented using their own lags. e_t - p_t is proxied by real consumption of nondurable goods. 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.

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