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*Monetary Disorder and Financial Regimes.
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Monetary Disorder and Financial Regimes. The Demand for Money in Argentina, 1900-2006*

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Abstract

Argentina is a unique experience of protracted economic instability and monetary disorder. In the framework of a long-term view, we investigate the demand for narrow money in Argentina from 1900 to 2006, shedding some light on the existence of money demand equilibria in extremely turbulent economies. The paper examines the effect of monetary regime changes by dealing with the presence of structural breaks in long-run equations. We estimate and test for regime changes through a sequential approach and we embed breaks in long-run models. A robust cointegration analysis can be hence performed in a single-equation framework. We find that estimated parameters are in sharp contrast with those reported in the literature for Argentina, but in line with those reported for industrialized countries, while significant structural breaks appear consistent with major policy shocks that took place in Argentina during the 20th century.

Keywords: Money demand, financial regimes, structural breaks, single-equation cointegration, cointegration test, Argentina monetary history.

JEL classification: E41, E42, C13, C22

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1 Introduction

Since the early days of nation building in the 19th century, Argentina has been a laboratory for monetary economics. A large variety of extreme economic episodes in the midst of political instability show in the history of the country: from rapid development in the *Belle Epoque* of 1880-1930 to economic crisis, from gold-standard stability to chronic monetary disorder, recurrent balance of payments crisis, speculative attacks on the currency, hyperinflation episodes and debt defaults (Diaz Alejandro, 1970; Véganzones and Winograd, 1997; Della Paolera and Taylor, 2001, 2003). The striking variety of macro policy experiments and monetary events made Argentina a fascinating and puzzling case-study for economic historians and applied economists.

The extreme monetary episodes observed in Argentina over the last 30 years, coupled with the recent advances in econometric theory, have led researchers to show increasing interest in the complex nature of these monetary phenomena. The research in modeling based on high-frequency data for short sample periods were used to disentangle the mechanisms driving money demand dynamics during the chronic high-inflation years of the 1970s and 1980s as well as the hyperinflation episodes of the late 1980s and early 1990s (Ahumada, 1992; Melnick, 1990; Kamin and Ericsson, 1993).

On the other hand, since the seminal works of Friedman (1956, 1959) and Friedman and Schwartz (1963, 1982) on the monetary history of the United States and the United Kingdom, low-frequency data for long sample periods have been at the heart of numerous studies in the money demand literature (Lucas, 1988, for the US; Hendry and Ericsson, 1991, for the US and UK; Muscatelli and Spinelli, 2000, for Italy). These works showed that historical data, along with simple econometric models, are able to fit the stylized monetary facts in most industrialized countries (Bordo and Jonung, 1987, 2003).

Accordingly, recent research has started to pay attention to the historical and econometric analysis of monetary aggregates in Argentina during the 20th century (Choudhry, 1995b; Urga and Winograd, 1997; Ahumada and Garegnani, 2002; Gay, 2004). However, these papers are often inconclusive on many issues raised by the existing theoretical and empirical literature on the long term view of monetary behavior. Such issues as the decision making model for rational asset allocation of economic agents under conditions of extreme volatility, as well as the existence of long-run monetary equilibria and the stability of money, require further research. The results obtained by the latter literature contradicts the general empirical evidence of stable relationships and parameter constancy well-established for most industrialized countries.

This paper proposes to address these issues for Argentina, using annual data from 1900 to 2006. Our empirical contribution focuses on the estimate of a long-run money demand model in a single-equation framework, providing economically interpretable opportunity cost parameters based on an econometric model with good statistical properties. We further consider the issue of shocks and long-run persistent financial and monetary distress in Argentina, as pointed out by Diaz Alejandro (1970) and Véganzones and Winograd (1997).

Drawing from the rich empirical literature on monetary instability, “missing money” and financial innovations in the industrialized countries (Judd and Scadding, 1982; Baba, Hendry and Starr, 1992; Carlson *et al.*, 2000), we account for financial and monetary (and, more generally, macroeconomic) regime changes. For this purpose, we date and test for structural breaks in the money demand function through a sequential bootstrap procedure. We consider M1 as the monetary measure in our money demand analysis.

From an empirical point of view, the behavior of the domestic currency allows us to explore a set of relevant economic perturbances in the monetary history of Argentina, such as currency substitution and speculative attacks, as well as persistent de-monetization and the chronic loss of confidence in paper money. In addition, a study on the narrow money aggregate allows us to “isolate the numerical *constants* of monetary behavior” (Friedman, 1956) and it would straightforwardly extend to further discussions about seignorage, the welfare cost of inflation and monetary hysteresis.

Our main findings are as follows. *First*, we deal with instability in long-run relationships obtained through a standard cointegration analysis. Structural breaks analysis suggests the presence of significant regime shifts in the money demand model, consistent with historically observed macroeconomic and financial shocks in Argentina. Stability and long-run equilibrium features are then crucially related to the estimate of relevant regime shifts. *Second*, unlike the existing literature on Argentina, we report evidence suggesting that the hypothesis of income homogeneity cannot be rejected in our money demand models. Estimated interest rate and inflation coefficients are both highly significant, while they sharply differ in magnitude. This may reflect the strong financial control and the credit market regulations that affected Argentina even during periods of high inflationary pressures, leading to persistent negative real interest rate on regulated deposits. *Third*, we explore the econometric issue of long-run equilibrium with regime changes, by implementing *ad hoc* tests for cointegration with multiple breaks. Results suggest that estimated long-run equilibria are consistent with the cointegration hypothesis. *Finally*, short-run estimates for money demand dynamics are reported. All in all, our simple model appears to fit the data in a robust way.

The remainder of this paper is as follows. In Section 2, we report a brief introduction to the monetary history of Argentina during the 20th century and early 21st century. A brief review of the theoretical and empirical literature is reported in Section 3. Section 4 presents the dataset and we show preliminary empirical results on unit root tests and cointegration: instability features arise in long-run equilibria when potential structural breaks for monetary regime changes are neglected. In Section 5, we date and test for structural breaks in the money demand model and we discuss their economic and historical relevance. Finally, in Sections 6 and 7 we report estimates for the money demand models. Section 8 concludes.

2 A Macro-Monetary Overview of Argentina, 1900-2006

2.1 The “golden age” period (1880-1930)

The development of the financial sector in Argentina dates back to the second half of the 19th century. This period shows an unprecedented economic effervescence, the so-called *Belle Epoque* (Diaz Alejandro, 1970; Véganzones and Winograd, 1997; Della Paolera and Taylor, 2001, 2003). Nonetheless, many attempts to anchor the monetary system to the international gold-standard failed due mainly to inconsistent macroeconomic policies, or the result of the international business cycle, leading to banking and financial crises, speculative attacks over the paper-gold peso parity and sharp devaluations of the currency. The “Baring crisis” due to the collapse of the public finances of Argentina in 1890 shocked the London financial market. The gold-standard regime was restored in 1899, becoming fully operative in 1900. In the framework of an economy fully integrated to international goods and capital markets, monetary policy in Argentina was then con-

ducted in accordance with the strict rule of the gold-standard regime. A Conversion Office (Caja de Conversion) and a national commercial bank (Banco de la Nacion Argentina) operated as independent monetary authorities, assuring a decade of remarkable financial stability until 1913.

Argentina was then and in the following 20 years in the top league of the ten richest nations of the time. The First World War led to the collapse of the gold-standard system, but the remarkable dynamism of the economy was still the source of the high level of development that Argentina had achieved. Nevertheless, the financial collapse of the United Kingdom in 1921, the short-lived experience of resumed convertibility in 1928-1929 and the incoming world economic crisis led to the end of the “golden age” (Diaz Alejandro, 1970).

2.2 The crisis of the 1930s and the politics of limited public intervention

The world crisis of 1930 broke the long phase of rapid growth and virtually cut off the inflow of foreign capital to Argentina. At the same time, the rapidly deteriorating terms of trade and the collapse of economic activity led to a series of bankruptcies, which further worsened the liquidity crisis in the banking system, turning it into a solvency crisis. The balance of payments crisis and the stress in the financial system provoked a speculative attack on the currency. A vicious circle of financial contraction and declining output was in motion¹.

Financial intermediation fell sharply from levels of 60-70% of GDP for the M3 aggregate to 50% in the first years of the crisis. The disruption of the financial system led to the insolvency of a number of banks. The government progressively liquidated the insolvent financial firms and created a hospital public institution to manage and sell the assets of other banks in difficulties. To contend with this adverse economic scenario, the government created the Central Bank in 1935. Its objectives were to ensure monetary stability and smooth the output fluctuations, the Central Bank was given a monopoly of issuing bank notes and was assigned to the role of lender of last resort. Additionally, the Central Bank replaced the Conversion Office for exchange operations.

In this period 1935-1945, since the creation of the Central Bank, we can identify two phases of monetary policy. The first one (1935-1940) is characterized by price stability, with the Central Bank compensating the monetary impact of the balance of payment flows. Indeed, the newly established Central Bank operated in a very prudent manner, intervening regularly in the market to stabilize monetary aggregates in the face of macroeconomic fluctuations. A second phase (1940-1945), under the conditions imposed by the World War, showed a marked improvement in the terms of trade. A large surplus in the balance of payments, mainly determined by growing exports to the United States and the United Kingdom, as well as decreasing imports, was accompanied by the generalized expansion of means of payment. This set up showed in the inflationary pressures, initially well managed by the monetary authorities, that will later turn into a chronic feature of the Argentine economy. Nevertheless, during the period under con-

¹As in other events of crisis, the events of 1930 showed the asymmetry of the automatic adjustment mechanism of the gold standard. The dynamics of adjustment work properly in the phases of expansion, when the balance of payments is in surplus, reserve accumulation contributes to the expansion of credit and output growth. However, in the phase of capital outflows the reverse dynamics is not a smooth process: deposits and credit contract, output decreases and a liquidity crisis is in motion.

sideration, the monetary policy aimed to absorb the excess of base money creation, thus avoiding an over-expansion of the money stock (Diaz Alejandro, 1970), while imported inflation mainly drove the domestic inflation pattern. This new regime of *limited public intervention* was effective in the objective of smoothing output fluctuations and attaining a reasonable degree of monetary stability. However, financial savings were durably affected by the impact of the world crisis on capital markets and the fragility of the local financial system. Disintermediation continued to decrease with M3 to GDP falling to 40% of GDP in the early 1940s (Véganzones and Winograd, 1997), but the average inflation rate of the period 1900-1940 was a low 1.8%.

2.3 Financial repression, disintermediation and chronic inflation (1945-1973)

The institutional changes were more radical after the end of World War II, when the first Peron government (1946-1951) was elected. The public authorities decided to “nationalize” bank deposits in 1946. The compulsory reserve requirements were fixed at 100% and all deposits of the commercial banks were managed in the name of the Central Bank for a fix fee charged by the commercial banks. Furthermore, in the new mandate of the Central Bank real interest rates were considered to be a “strategic” policy variable and set at negative levels for almost all the period until the financial liberalization of 1977. In the Peron years of 1946-55, under a regime of *massive public intervention* in the goods and financial markets, expansionist monetary and fiscal policies were followed, and inflation (in the range of 4-40%) became a chronic feature of the monetary behavior of the economy.

A first step towards the liberalization of the financial system occurred in 1957 after the coup d'état that deposed General Peron in 1955. The nationalization of deposits was abandoned, the legal minimum reserves were reduced and the credit allocation policy once again decentralized. During the 1960s, the financial sector experienced the rapid development of a large number of specialized non-banking institutions. However, monetary financial instruments continued to fare badly due to inflation that reached a peak of 80% in 1959, as the result of sharp devaluations. By the end of 1960s the government tried to integrate the newly established non-banking institutions into the official system. But the appointment of an elected government headed by General Peron in 1973 set the beginning of a new regime, marked by accelerating inflation, macroeconomic instability and political turbulence, that will end again in a coup d'état and dictatorship in 1976..

2.4 High inflation, financial fragility and stabilization failures (1974-1990)

The newly elected Peronist government launched in 1973 a stabilization plan based on a rigid income policy coupled with expansionary fiscal and monetary policies. In 1975, after the death of Peron in 1974 and the extreme political disturbances that followed, a new stabilization plan collapsed on the attempt of a sharp readjustment of relative prices to contain macroeconomic disequilibria (the so-called *Rodrigazo*). This stabilization failure set the motion for a long lasting period of chronic high inflation, reaching levels of 80-600% a year and, finally the experience of hyperinflation in 1989 and 1990.

After the military coup of 1976, a financial reform was planned and executed in 1977,

when the financial system was liberalized and interest rates (for the first time) were freed. In addition, the reserve requirements were decreased (and remunerated by the Central Bank) and credit controls were further relaxed for private commercial banks. A second stabilization plan was launched in December 1978, the *Tablita Plan*, focusing mainly on a pre-announced rate of devaluation for a set period of time in view of the gradual reduction of inflation. But the absence of public finance consistency and persistent inflation over the targets anticipated by the authorities led to sharp appreciation of the currency, rising current account deficits and increasing public debt. These policies failed to stabilize the financial system and the Argentinean economy, leading to persistent high inflation, a confidence crisis and speculative attacks. The instability resulted in the banking (and more general macroeconomic) crisis of 1980. The *Tablita* plan collapsed in 1981, a sharp devaluation took place and by 1982 the country will face a disruptive debt crisis along other Latin American countries.

Following the banking crisis of the 1980s, the financial system was hardly altered. In 1983 with the return to democracy, after the Malvinas war of 1982, President Alfonsín (Radical Party) was elected, in the face of extremely high inflation and severe macroeconomic instability. The implementation of the *Austral Plan*, a monetary reform implemented in mid-1985 involving wages, prices, and exchange rate freezes, and later the *Spring Plan* (a mid-1988 agreement with the private sector to limit the growth of prices and the exchange rate devaluation at 4% per month) failed to deal with exploding inflation (Kamin and Ericsson, 2003). The collapse of public finance, rapidly rising political instability and recurrent speculative attacks on the currency led to the hyperinflationary episodes of 1989 and 1990.

The degree of demonetization then reached extreme levels, with the M1 to GDP ratio standing at 1.5% in 1990 and M3 to GDP reaching low levels of 5%, compared to 70% in the first decades of the century. Monetary stabilization was obtained in 1991, under President Menem (Peronist Party) elected in 1989, with the *Convertibility Plan* (March 1991), which fixed the parity against the dollar and led to a sharp reduction of inflation to one digit yearly levels.

2.5 The Convertibility Law of 1991: successful stabilization and monetary hysteresis (1991-2000)

The Convertibility Law of 1991 and the creation of a Currency Board regime were introduced to address the credibility problem of chronic monetary disorder and the resulting massive dollarization of Argentina's financial instruments. The system thus consisted of a simple dollar standard put into practice as a bi-monetary system with both the dollar and peso as legal tender. The monetary system was rather similar to the gold-standard rule at the beginning of the century. Furthermore, the new law banned the Central Bank from monetary financing of the public debt, proclaiming *de facto* its independence, later *de jure*. The creation of money was only allowed simultaneously with increases in foreign currency reserves. Finally, the role of the Central Bank as lender of last resort was restricted. The monetary initiative was coupled with an intensive programme of privatization and regulatory reforms.

The reduction of inflation was sharp and rapid, as in other post hyperinflation episodes. In the midst of the high liquidity of international capital markets, the newly gained confidence of Argentina led to massive capital inflows and solid reserves accumulation. This provided a basis for the monetary recovery of the country, but the pattern

was probably mined by a hysteresis phenomenon affecting the amount of liquidity held by agents and recovery resulted quite slow (the M1/GDP ratio rose to 4.4% in 1992, but it only attained 6.8% in 1999). However, the sudden disinflation led to a consumption boom, reinforced by capital inflows and public finance disequilibria. A gradual currency appreciation and increasing current account deficits cum growing indebtedness will eventually set the conditions for a credibility problem on the stability of the peso parity and the convertibility regime. The debt crisis in Mexico (1994) and later the emerging market crisis in Asia, Russia and in neighboring Brazil, with the contagion effects, only reinforced the confidence crisis on the sustainability of the monetary regime of Argentina. The fundamental conditions for the speculative attacks and deep economic crisis of 2001-2002 were set.

2.6 The end of Convertibility, debt default and back... (2001-2006)

Fiscal deficits, international financial crises and overvaluation of the peso reverted the virtuous circle of the early nineties. This generated pressures over commercial banks' liquidity and interbanking rates, spreading to the whole financial system and leading to a bank panic and a bank run over the domestic commercial deposits in 2001. The convertibility system was kept unchanged until January 2002, when the parity with the dollar was abandoned. The peso suffered a devaluation of 300% and, consequently, inflation skyrocketed to 40% yearly in the midst of a severe political crisis that led to the resignation of President De la Rúa and a rapid succession of transitory administrations in a few weeks.

Economic policy turned into a rather chaotic exercise, with deposit freeze, a monetary reform, banking rescue packages and debt default, amid daily demonstrations of the population in a state of shock and deep social distress. The Central Bank played a crucial role in the stabilization of both the exchange rate and the inflation. In particular, this was achieved through a mix of interventions in the domestic money market and in the foreign exchange market. After a deep contraction of output of 6% in 2002, the economy started a rapid and solid recovery in 2003 sustained by very favorable external conditions of strong improvement of the terms of trade, and a recovery of public finance. Inflation fell to one digit number to rise thereafter to 8-20% range in the coming years until the end of the first decade of the 21st century. Since the recovery, real narrow aggregates grew at faster rate, with the M1/GDP ratio attaining 12.3% in 2006, in a rather stable economic environment. However, we should highlight that, post hyperinflation cum price stabilization in 1991 and later, financial disintermediation showed strong signs of persistence and hysteresis. Argentina's financial system never returned to the high values of the *golden age*.

3 Money demand in Argentina

There is a fairly large body of research on the theoretical structure of money demand. Judd and Scadding (1982), Goldfeld and Sichel (1990) and Sriram (1999) offer, among others, an exhaustive review on the theoretical and empirical literature.

From a theoretical point of view, political economy models offer a solid support for the analysis of monetary behaviors dealing with protracted inflation regimes, hyperinflation diseases, financial adaptations (currency substitution, capital flights, demonetiza-

tion) and stabilization attempts². The political economy framework fits particularly well in the case of Argentina, where inflation processes and distributive struggle, financial instabilities and monetary regimes are main long-term policy elements.

Despite recent advances introduced in the monetary theory³, the money demand function used in empirical applications is common to most models developed in the literature since the seminal works of Keynes (1930, 1936), Baumol (1952), Tobin (1956, 1958) and Friedman (1956, 1959). A standard money demand function can be formalized as:

$$M = f(Y, P, \mathbf{R}) \tag{1}$$

where M is the nominal stock of money, P is the prices level, Y is the scale variable (*i.e.* an income variable, usually defined in real terms) and \mathbf{R} is a vector of opportunity costs in holding nominal balances. For simplicity, the function $f(\cdot)$ is assumed to be unit homogeneous in prices, increasing in income and decreasing (increasing) in the vector of the opportunity costs which reference assets are outside (inside) the monetary aggregate M . The price homogeneity hypothesis insures that the money demand function can be reformulated in terms of real aggregates, rather than nominal, without loss of generality.

Many attempts to estimate a money demand function for Argentina can be found in the literature. Most empirical studies pay more attention to the high-inflation or hyperinflation episodes of the 1980s and to the monetary effects of the Convertibility plan during the 1990s, by using high-frequency short samples. Ahumada (1992) and Melnick (1990) develop models of real M1 demand in Argentina between 1977 and 1988, showing the existence of non-linearities in the short-run dynamics of money demand caused by high and rising inflation rates. Choudhry (1995a) restrains the analysis on the existence of stationary long-run M1 and M2 functions between 1975 and 1987, finding significant results for both monetary aggregates. The monetary biases (non-linearities) generated by the hyperinflation episode are mainly handled by Kamin and Ericsson (1993), who develop a model of real M3 demand between 1978 and 1993. Finally, Aguirre *et al.* (2006) analyze the demand for money during the Convertibility period, the economic crisis and the recovery (1993-2005), finding poor results in terms of stability of the implied monetary relationships.

In sharp contrast, only a few empirical studies on low-frequency data deal with the issue of the long-run stability of money demand in Argentina. Choudhry (1995b) find a stationary long-run money demand relationship between 1935 and 1962, using quarterly values for real M1 and M2, real net national income and an inflation rate. However, results inconsistent with theoretical priors may suggest some form of misspecification in the underlying VECM. Gay (2004) analyze the demand for real M1 balances in the context of a small open economy model with nontradable goods. Evidence of endogeneity in the income variable may suggestively point out the non-neutrality of money in Argentina or, in turn, some misspecifications in the long-run equation. Finally, Urga and

²These models develop a distributive conflict framework with heterogeneous agents and optimal inflation tax (Laban and Sturzenegger, 1994; Mondino *et al.*, 1996; Ahrend, Verdier and Winograd, 2008).

³See, for instance, the general transaction and portfolio joint models (Ando and Shell, 1975; McCallum and Goodfriend, 1987) and models with money in a general equilibrium framework, as the models with imperfect markets (Bewley, 1983; Kehoe *et al.*, 1990) and the class of search models (Kiyotaki and Wright, 1989; Trejos and Wright, 1995).

Winograd (1997) and Ahumada and Garegnani (2002) deal with the long-run instability issue of the money demand relationship in Argentina. The latter use annual data for real M1 from 1935 to 2000 and introduce exogenous breaks in the long-run parameters⁴, while the former show that the demand for real M1 between 1900 and 1992 is affected by few (endogenously estimated) structural breaks.

Results reported in previous works highlight the crucial issue of the long-run instability of the money demand in Argentina. The presence of abrupt changes in monetary regimes, as suggested by the evident statistical difficulty in describing monetary phenomena, may jeopardize the standard econometric approach. In this paper we address the issue of modeling instability making use of recent developments in econometric treatment of breaks.

4 Preliminary Empirical Analysis

4.1 Data

Our dataset spans from 1900 to 2006, *i.e.* from the beginning of the long lasting gold-standard experiment to the aftermath of the deepest economic crisis ever experienced in Argentina (See the Appendix for more details). Let m be the log of the nominal M1, *i.e.* the sum of the total currency held by the public, plus small demand deposits. The annual (log) real money aggregate ($m-p$) is computed as the annual average of monthly nominal values of M1, divided by the respective monthly value of the wholesales price index. This allows us to assign a weight to large deviations of prices and to better account for the infra-annual volatility of monthly real balances. The transaction variable is the (log) real GDP (y), computed as the annual nominal GDP divided by its implicit price deflator. The price index chosen for this study is the wholesales price index (p), rather than the CPI series, because of data quality and availability. As before, the series is computed as an annual average of monthly price index values. The inflation rate (π) is thus computed as the first difference of the annual log price series. Finally, the external opportunity cost in holding real balances is the annualized interest rate over relevant alternative assets (r)⁵. In Figure 1 we plot the series, scaled in log values.

[Figure 1 about here]

4.2 Unit-root tests

The very long time span of our dataset and the high degree of volatility should allow for some non-stationarity in mean and variance of the series. After a visual inspection,

⁴The income homogeneity restriction ($\beta_y = 1$) is assumed to hold from 1935 to 1955, while the Baumol-Tobin restriction ($\beta_y = 0.5$) is assumed to hold from 1956 to 2000. In addition, the interest rate is assumed to be the relevant opportunity cost from 1935 to 1973 and from 1991 through the end of sample, while the inflation is the relevant opportunity cost from 1973 to 1991.

⁵Given the hard task to rebuild homogeneous series for interest rates throughout the century, we used different measures for the external opportunity cost. Rates over time deposits and saving deposits, as well as regulated and non-regulated interest rates and free market rates, were then alternated to recompose a relevant interest rate series. The series obtained in a such way enters the model as $\ln(1+r_t)$, which is a tantamount to the semi-log specification. Indeed, a double-log specification setting would not be coherent with the extreme variation of nominal variables (interest rate and inflation rate) during high-inflation episodes.

we cannot reject this prior. We check this hypothesis using a battery of efficient unit root tests introduced by Ng and Perron (2001), such as the modified versions of the Elliott, Rothenberg and Stock (1996) feasible point-optimal test (MP_T^{GLS}) and GLS Dickey-Fuller test (ADF^{GLS}), as well as the GLS Phillips-Perron t -test (MZ_T^{GLS}). The power of this class of modified unit root tests is dramatically improved by the use of the GLS detrending of the series. The optimal lag selection criterion for the computation of the Ng and Perron (2001) autoregressive spectral density function is corrected following Perron and Qu (2007).

Table 1: Unit-root tests

Series	ADF^{GLS}	MZ_T^{GLS}	5% cv	MP_T^{GLS}	5% cv
$(m-p)^+$	-1.96	-1.92	(-2.91)	12.09	(5.48)
y^+	-1.71	-1.60	(-2.91)	15.65	(5.48)
π	-2.04**	-1.93	(-1.98)	3.28	(3.17)
r	-1.60	-1.55	(-1.98)	5.12	(3.17)

Note: $^+$ a constant and a linear trend are included into the test equation. 5% critical values for ADF^{GLS} and MZ_T^{GLS} tests are from Fuller (1976), whereas for the MP_T^{GLS} test 5% critical values are from Ng and Perron (2001). ***, ** and * denote significance at 1, 5 and 10%, respectively.

On the basis of results reported in Table 1, the null hypothesis of unit root cannot be rejected for all variables, even though the ADF^{GLS} and the MZ_T^{GLS} statistics are very close to the 5% rejection area for the inflation series. Our results do not conflict with findings reported elsewhere (for instances, Ahumada and Garegnani (2002) and Choudhry (1995b)) when different sample windows are used.

4.3 Cointegration analysis

We employ cointegration techniques to estimate long-run money demand parameters. The multivariate system considered here involves the monetary aggregate, one transaction variable and two measures of the opportunity cost in holding real balances.

The macroeconomic theory on open economies asserts the existence of a theoretical equilibrium between prices, interest rates and the exchange rate. In the long-run, this means that the PPP (Purchase Power Parity) and the CIRP (Covered Interest Rate Parity) should hold. The international nominal equilibria lead to a domestic equilibrium between nominal interest rates and the (expected) change in prices, that is a Fisher parity in the domestic real interest rate.

Assuming that in the very long-run the international equilibria hold, the dimension of the cointegration space must be characterized by the existence of at least two theoretical long-run relationships. Thus, a multivariate cointegrated system involving real money balances, real transactions, the inflation rate and the nominal interest rate must show a rank of order 2. Thus, the first cointegrating vector is the money demand equation and the second cointegration relationship is the real interest rate equation. However, in order to identify individual cointegrating vectors, we must impose an exclusion restriction over at least one variable for each relationship⁶.

⁶We exclude the exchange rate from the long-run analysis for two reasons. It should be an important

This means that we can identify a money demand model involving opportunity cost variables one by one, *i.e.* a model with the inflation rate and a model with the nominal interest rate. If the strong Fisher parity hypothesis is supported by data, *i.e.* the homogeneity restriction $\{1, -1\}$ over coefficients is not rejected, the two models are statistically equivalent and it is irrelevant which opportunity cost is excluded from the equilibrium specification (Hoffman *et al.*, 1995). Otherwise, the two models are not equivalent and a different inference would be necessary.

Cointegration between non-stationary variables is tested through the Stock and Watson (1988) test for common stochastic trends and the Johansen (1988, 1991) trace and maximum eigenvalue tests. Results for the real interest rate equation and for the two money demand models described above are reported in Table 2.

The upper part of the table reports cointegration tests. According to results, the null hypothesis of no cointegration can be strongly rejected for the real interest rate equation, as well as for the money demand model. In addition, the null hypothesis of a single cointegrating vector can also be rejected for the money demand relationship, confirming the existence of two long-run relationships. Further, we estimate the rank 2 model by maximum likelihood, imposing an exclusion restriction over either the inflation rate (Model 1) or the interest rate (Model 2). The first cointegrating equation is then the money demand model, while the second cointegrating equation describes the real interest rate relationship. Restrictions over long-run coefficients are tested and imposed. In particular, the Baumol-Tobin restriction over the real GDP parameter ($\beta_y = 0.5$) cannot be rejected for both models 1 and 2, while the strong Fisher parity hypothesis must be rejected⁷. For our purposes, the latter result implies that the two money demand models are not statistically equivalent and two different modeling strategies would be necessary.

Further, we check the stability of the estimated money demand cointegrating relationships. We consider two statistics developed by Hansen (1992) and based on the Phillips and Hansen (1990) Fully-Modified estimator. The first statistic tests for the null hypothesis of stable cointegration against the alternative of an unknown single change in the deterministic and stochastic long-run parameters (*SupF*), while the second one assumes a constant hazard of parameters instability (*MeanF*) as the alternative hypothesis. Results are reported in Table 3. Test statistics strongly reject the null hypothesis of stable cointegration for both Models 1 and 2. It is worth noticing that the statistical instability of cointegrating vectors is contradictory with the idea of cointegration. Quite surprisingly, our results suggest that cointegration tests spuriously detect cointegration when the underline model neglects some unknown feature of data acting as a source of long-run instability.

element if we were to test for a long-run PPP and CIRP equilibria, but in this analysis we suppose that these classical international conditions hold. Second, changes in the exchange rate are a key element in the currency substitution theory and unofficial dollarization was an important issue in Argentina during the high-inflationary episodes in 1970s and 1980s. However, given the existence of wide periods of fixed exchange rate regimes (convertibility under the gold-standard and the 1991 stabilization plan), we assume that prices and interest rates embed enough information in the long-run to capture most dynamics of the demand for real balances.

⁷Table 2 reports results involving a large set of significant restrictions on long-run coefficients matrix β and adjustment coefficients matrix α . When the quantitative theory restriction on the scale variable coefficient ($\beta_y = 1$) is imposed, the LR statistic rejects both models 1 and 2, at less than 1%. In addition, the adjustment coefficient $\alpha_{(m-p)}$ is not significant. When the strong Fisher parity restriction

Table 2: Cointegration tests

	Real Interest Rate (r, π)			Money Demand ($m - p, y, r, \pi$)		
	Statistic	5% cv	p -value	Statistic	5% cv	p -value
$J^{Trace}(0)$	27.9	20.3	< 0.01	62.4	47.8	0.01
$J_{\mu}^{Max\lambda}(0)$	22.1	15.9	< 0.01	29.1	27.6	0.03
$Q_{\mu}^f(0)$	-68.2	-30.9	< 0.01	-59.9	-39.3	< 0.01
$J^{Trace}(1)$	-	-	-	33.4	29.8	0.02
$J_{\mu}^{Max\lambda}(1)$	-	-	-	25.4	21.1	0.01
$Q_{\mu}^f(1)$	-	-	-	-59.9	-22.5	< 0.01

	Money Demand (Model 1)		Money Demand (Model 2)	
	CE 1	CE 2	CE 1	CE 2
$(m - p)$	1	-	$(m - p)$	1
y	-0.5	-	y	-0.5
r	2.0***	1	r	-
π	-	-0.6***	π	1.16***

LR Test for Restrictions		LR Test for Restrictions	
$\chi^2(3)$	5.38	$\chi^2(4)$	5.66
p -value	0.14	p -value	0.22

Note: J^{Trace} and $J^{Max\lambda}$ are the Johansen (1988, 1991) trace and maximum eigenvalue tests, respectively. 2 lags of the endogenous variables are included in the level VAR, according to the Hannan-Quinn information criterion, for all models tested. MacKinnon, Haug and Michelis (1999) p -values are computed. The models include a constant in the cointegration equation and into the VAR. Q^f is the Stock and Watson (1988) common trend statistic. A constant term is included into the test equations. In parantheses, the number of cointegrating vectors under the null hypothesis. Maximum likelihood estimates for cointegrating models and likelihood ratio tests for binding restrictions are reported in lower panels.

***, ** and * denote significance at 1, 5 and 10%, respectively.

The monetary history of Argentina can help us to shed some light on this issue. As already mentioned in Section 2, Argentina went through many political, economic and monetary regimes since the beginning of the 20th century. This, we can expect that the switch from one regime to an other is the main source of instability for money demand relationships. For instance, a jump in monetary equilibria are a natural issue for a country either facing the collapse of a monetary system or dealing with rapid institutional changes. On the other hand, new monetary regimes can be set by long lasting persistent features affecting the demand for money in the long-run, as well as hysteresis effects over monetary behaviors and financial innovations/substitutions. According to this definition of regime change, our long span data analysis is obviously affected by the extreme economic features which took place in Argentina during the century, such as abrupt adjustments after full convertibility regimes and persistent money holding behaviors driven by chronic inflationary pressures. Neglecting these features would mean disregarding the stylized facts and reporting inaccurate econometric results.

($\beta_{\pi} = -\beta_r$) is imposed, the LR statistic rejects again both models 1 and 2, at 5% and 1% respectively.

Table 3: Test for parameters instability

Statistic	Value	<i>p</i> -value
Money Demand Model 1 ($m - p, y, r$)		
<i>MeanF</i>	12.0	< 0.01
<i>SupF</i>	23.4	< 0.01
Money Demand Model 2 ($m - p, y, \pi$)		
<i>MeanF</i>	14.4	< 0.01
<i>SupF</i>	26.3	< 0.01

Note: Test statistics are based on the Fully-Modified estimator (Phillips and Hansen, 1990). The test equation includes a constant. The long-run covariance matrix is estimated through a Quadratic Spectral kernel and an automatic bandwidth selection (Andrews, 1991). The selected bandwidth is 9.6 for Model 1 and 14.5 for Model 2.

We propose to embed monetary regime changes in our models by estimating structural breaks in the money demand function and then including them in the long-run equations. After a visual inspection of the velocity of money in Figure 2, we claim that our definition of regime changes motivates the choice of deterministic structural breaks, rather than stochastic breaks. This approach should allow us capture variations in financial equilibria, monetary trends and some institutional and economic unobserved factors, which may probably explain the instability of estimated cointegrating equations⁸.

[Figure 2 about here]

5 Looking for structural breaks : the sequential estimation approach

We implement the sequential estimator developed by de Peretti and Urga (2004) for dating and testing break points in a multivariate framework. Following this approach, a vector of structural breaks is firstly estimated on the marginal process of a given system, and then this information is incorporated in the conditional process in order to estimate its own structural breaks.

Consider a bivariate model $w_t = (z_t, x_t)'$ with joint density function $f_w(w_t, \theta) = f_{z|x}(z_t|x_t, \lambda_1) \cdot f_x(x_t, \lambda_2)$, where the first term in the right hand side of the equation is the conditional density of z_t given x_t and the second term is the marginal density of x_t . Following de Peretti and Urga (2004) and Banerjee, Lazarova and Urga (1998), the reduced form system used to dating breaks is :

⁸In other words, we depart from a simple linear specification. The aim is to replicate the actual dynamic of the monetary phenomena under analysis, which is probably characterized by small and big shocks. By handling structural breaks, we are more confident about picking up observable big shocks rather than small ones. Anyway, while the former are relevant in our long span analysis, the latter would add relatively poor information.

$$\begin{aligned}
z_t = & \mu_0 + \delta_0 t + \sum_{\tau=1}^{p_0} \rho_{0,\tau} z_{t-\tau} + \sum_{i=1}^{I_0} (\alpha_{0,i} B_{0,i,t}^M + \beta_{0,i} B_{0,i,t}^L) + \\
& + \gamma x_{t-1} + \sum_{i=1}^{I_n} (\eta_i B_{1,i,t}^M + \theta_i B_{1,i,t}^L) + \varepsilon_t
\end{aligned} \tag{2}$$

$$x_t = \mu_1 + \delta_1 t + \sum_{\tau=1}^{p_1} \rho_{1,\tau} x_{t-\tau} + \sum_{i=1}^{I_1} (\alpha_{1,i} B_{1,i,t}^M + \beta_{1,i} B_{1,i,t}^L) + \nu_{n,t} \tag{3}$$

for $t = 1, 2, \dots, T$, where Equation (2) is the conditional process and Equation (3) is the marginal process. Dummy variables $B_{i,t}^M = \mathbf{I}(t \geq b_i)$ and $B_{i,t}^L = (t - b_i + 1)\mathbf{I}(t \geq b_i)$ are designed to capture breaks b_i in the mean and in the linear trend, respectively. Marginal and conditional processes also include p lags of the dependent variable.

Breaks dating is obtained through a sequential estimation which includes the breaks one by one. The structural breaks estimator is based on the conventional F -statistic estimator (Christiano, 1992), *i.e.* \hat{b}_i is defined as the time t at which the F -statistic reaches its maximum ($\hat{b}_i = \operatorname{argmax}\{F_t\}$). However, the sequential estimate of the structural breaks is showed to be biased in the presence of more than one break: the dating of the first break is biased when a second break is neglected, and so on. To overcome this problem, a backward revision of estimated breaks is proposed.

The sequential estimation approach usually poses the problem of stopping the search when further estimated breaks are no longer significant. The conventional asymptotic test for this hypothesis is showed to lead to an important bias in the rejection of structural stability (Christiano, 1992). Asymptotic alternatives proposed by Andrews (1993) and Andrews and Ploberger (1994) cannot be employed in the case of time series with unit roots and/or for systems with one or more structural breaks in the marginal process (Hansen, 2000). A bootstrap approach to test for the null hypothesis of $i - 1$ breaks against i breaks is here employed. We then stop the search when two consecutive estimated breaks are insignificant.

First, Equation (3) is estimated under the null hypothesis of $i - 1$ breaks. The conventional F -statistic estimator is used to estimate breaks under the null and the statistic $\hat{\tau} = \max\{F_t\}$ is computed. Bootstrap errors are then generated (parametrically or non-parametrically), and the estimated parameters together with bootstrap errors are used to generate a recursive bootstrap sample. The resampling procedure is repeated B times and for each iteration $j = 1, \dots, B$ the statistic $\tau_j^* = \max\{F_{t,j}\}$ is computed. The feasible bootstrap p -value is computed as:

$$p^*(\hat{\tau}) = \frac{1}{B} \sum_{j=1}^B I((\tau_j^*)^2 > (\hat{\tau})^2) \tag{4}$$

Finally, significant breaks are included in Equation (2), together with explanatory variables, and the sequential search is reiterated for the conditional process.

From Equation (3), a preliminary analysis involves the proper selection of the functional form of the data generating process (DGP) depicting the series. Following Chris-

tiano (1992), we investigate the $AR(p)$ form of the DGP for each marginal process under the null hypothesis of no breaks. Models are selected based on the analysis of the $AR(p)$ fitted residuals. Results (not reported) suggest the use of $p = 4$ lags for both the inflation and the interest rate models and $p = 2$ lags for the real GDP model⁹.

We then estimate and test for structural breaks in marginal and conditional models through the bootstrap approach. For the former, each equation includes a constant, a trend and p lags of the dependent variable. For the latter, models are estimated as displayed in Equation (2), *i.e.* replacing regressors with their $AR(p)$ marginal processes and their own structural breaks ($B_{n,i,t}^M$ and $B_{n,i,t}^L$). The lagged-dependent variable does not enter the equation, to avoid the problem of unit roots and insignificance of regressors when the regressand is non-stationary. Results are reported in Table 4 and 5.

Bootstrap p -values for the null hypothesis of $i - 1$ against i breaks are computed through $B = 999$ random replications. Three bootstrap sampling methods are implemented in this paper: parametric, non-parametric with d.o.f. correction and non-parametric with the correction advocated by Weber (1984).

5.1 Breaks in the marginal processes

5.1.1 Interest rate

Most significant breaks in the interest rate cover the 1980s, a period marked by a high fiscal deficit and the end of the military government in 1983, many inflation and interest rates peaks (1981, 1985 and 1987) and the beginning of the hyperinflation episode in 1989. Two more significant breaks are estimated in 1976 and 1978, which coincide with the end of the third Peronist government and the military coup (1976) and the effects of the strong financial reform implemented in 1977 by authorities. Although the procedure is not able to detect accurately this financial liberalization, the effects of another financial reform are captured by a significant break in 1957. Finally, the two breaks in 1992 and 2003 are coherent with the full recover of Argentina in the aftermath of the hyperinflation and of the economic crisis (2001-2002).

5.1.2 Inflation

Two breaks are jointly significant under both the parametric and the non-parametric bootstrap. The first one (1989) coincides again with the beginning of the hyperinflation episode spanning from 1989 through 1990. The second one (1975) coincides with the abrupt adjustment in prices and tariffs set by the Minister of the Economy, Celestino Rodrigo, through a strong exchange rate devaluation. This adjustment (called *Rodrigazo*) was decided to fix the distortion of relative prices, due to the regime of price controls in force since 1973, but caused a boom of inflation (followed by trade union strikes and slowdown of economic activity) with persistent effects until the 80s (monthly wholesales inflation registered a jump from 5% up to 44% between May and June 1975, when the adjustment was implemented). Another break is detected in 1996, but it cannot be rejected only by the parametric bootstrap-based test.

⁹Since all variables are non-stationary, autoregressive models are estimated in first differences. In addition, models showing a drift in levels are estimated without a constant in first differences, while trending variable regressions are estimated with a constant. Optimal lag-length selection for $AR(p)$ models is based on the Ljung-Box test statistic and the Schwarz information criterion jointly. Results are available upon request.

Table 4: Structural breaks in marginal models

significant breaks							
r_t	(i)	(ii)	(iii)	(iv)	(v)	(vi)	(vii)
Date	1989	1987	1976	1983	1981	2003	1985
PB p -value	0.00	0.00	0.00	0.00	0.00	0.00	0.00
NPB(1) p -value	0.00	0.00	0.00	0.01	0.04	0.01	0.01
NPB(2) p -value	0.00	0.00	0.01	0.01	0.02	0.00	0.01

significant breaks							
r_t	...	(viii)	(ix)	(x)	(xi)	(xii)	(xiii)
Date	...	1992	1957	2001	1978	1933	1974
PB p -value	...	0.00	0.01	0.27	0.00	0.21	0.47
NPB(1) p -value	...	0.00	0.02	0.63	0.00	0.21	0.51
NPB(2) p -value	...	0.00	0.02	0.73	0.00	0.22	0.48

significant breaks					
π_t	(i)	(ii)	(iii)	(iv)	(v)
Date	1996	1989	1975	1986	1984
PB p -value	0.02	0.00	0.00	0.07	0.06
NPB(1) p -value	0.13	0.00	0.00	0.18	0.14
NPB(2) p -value	0.14	0.00	0.00	0.13	0.13

y_t	(i)	(ii)	(iii)	(iv)	(v)	(vi)	(vii)
Date	1964	1913	1918	1977	1931	2002	1989
PB p -value	0.22	0.88	0.54	0.35	0.21	0.07	0.03
NPB(1) p -value	0.24	0.89	0.54	0.31	0.25	0.12	0.04
NPB(2) p -value	0.24	0.88	0.54	0.36	0.24	0.11	0.04

Note: PB is the parametric bootstrap; NPB(1) is the non-parametric bootstrap with d.o.f. correction; NPB(2) is the non-parametric bootstrap with the correction proposed by Weber (1984). Bootstrap samples are generated through 999 random replications. The order of the estimated breaks is sorted as in outcome of the sequential procedure.

5.1.3 Real income

Although we have strong priors about regime changes in the real economic activity of Argentina, the sequential procedure is not able to identify any significant structural break even after many search stages. We can only find a significant break in 1989, *i.e.* the beginning of the hyperinflation episode. Using different model specifications did not improve results. We therefore accept the hypothesis of one single structural break.

5.2 Breaks in the money demand models

Given the set of estimated breaks in marginal models, test equations for conditional models take the following form:

$$\begin{aligned}
(m-p)_t = & \mu_0 + \delta_0 t + \sum_{i=1}^{I_0} (\alpha_{0,i} B_{m-p,i,t}^M + \beta_{0,i} B_{m-p,i,t}^L) + \sum_{j=1}^2 \gamma_{1,j} y_{t-j} + \\
& (\eta_{1,1} B_{y,1,t}^M + \theta_{1,1} B_{y,1,t}^L) + \sum_{j=1}^4 \gamma_{2,j} r_{t-j} + \sum_{i=1}^{10} (\eta_{2,i} B_{r,i,t}^M + \theta_{2,i} B_{r,i,t}^L) + \varepsilon_t
\end{aligned} \tag{5}$$

$$\begin{aligned}
(m-p)_t = & \mu_0 + \delta_0 t + \sum_{i=1}^{I_0} (\alpha_{0,i} B_{m-p,i,t}^M + \beta_{0,i} B_{m-p,i,t}^L) + \sum_{j=1}^2 \gamma_{1,j} y_{t-j} + \\
& (\eta_{1,1} B_{y,1,t}^M + \theta_{1,1} B_{y,1,t}^L) + \sum_{j=1}^4 \gamma_{2,j} \pi_{t-j} + \sum_{i=1}^2 (\eta_{2,i} B_{\pi,i,t}^M + \theta_{2,i} B_{\pi,i,t}^L) + \varepsilon_t
\end{aligned} \tag{6}$$

Estimated breaks $B_{m-p,i,t}^M$ and $B_{m-p,i,t}^L$ are money demand's own breaks. Results are reported in Table 5.

Table 5: Structural breaks in conditional models

	significant breaks						
$m-p, y, r$	(i)	(ii)	(iii)	(iv)	(v)		
Date	1946	1959	1913	1974	1972		
PB p -value	0.00	0.05	0.06	0.71	0.56		
NPB(1) p -value	0.00	0.05	0.06	0.77	0.62		
NPB(2) p -value	0.00	0.07	0.06	0.77	0.57		

	significant breaks						
$m-p, y, \pi$	(i)	(ii)	(iii)	(iv)	(v)	(vi)	(vii)
Date	1959	1946	1986	1913	1997 ¹	1972	2003
PB p -value	0.00	0.00	0.01	0.03	0.00	0.28	0.14
NPB(1) p -value	0.00	0.00	0.02	0.02	0.01	0.29	0.16
NPB(2) p -value	0.00	0.00	0.02	0.02	0.01	0.28	0.14

Note: See Table 4. ¹ the break is detected when the 1996 date is excluded from the marginal process of inflation.

The upper panel reports results for the money demand model with the interest rate as opportunity cost. We have evidence of 3 structural breaks, in 1913, 1946 and 1959¹⁰. The first break (1913) coincides with the end of the gold-standard experience

¹⁰Since some instability affected the estimate of the first relevant break in the output, we obtained the final sequence of breaks through the following procedure: *first*, significant breaks in 1959 and 1913 were detected; *second*, these breaks were considered as known and the first relevant break was estimated again; *finally*, the resulting relevant break (1946) was considered as known while relaxing the hypothesis over breaks in 1959 and 1913.

of Argentina, a relatively stable economic and monetary regime since the beginning of the century, that is the zenith of the so-called *Belle Epoque* (Diaz Alejandro, 1970; Véganzones and Winograd, 1997; Della Paolera and Taylor, 2001, 2003). With respect to the second break (1946), the market and trade oriented economy and the financial development of the country were progressively reverted by nationalizations policies during the first Peron's mandate (1946-1951). High fiscal deficits, strong controls of the banking sector and inflationary pressures started to deteriorate the credibility in the financial sector, causing a disintermediation phenomenon which affected Argentina for most the second half of the century. The third break (1959) coincides with implementation of the Alsogaray's stabilization and development plan (July 1959), which aim was to raise economic growth (mainly through incentives to foreign investments and import substitution) and to stabilize prices (by fixing wages, prices and tariffs). The program took place during the mandate of President Frondizi (1958-1961) as a response to the explosion of prices due to the second highest peak in the fiscal deficit of the central government ever experienced in Argentina (Della Paolera and Taylor, 2003).

The lower panel in Table 5 reports results for the money demand model with the inflation as opportunity cost. In addition to breaks in 1913, 1946, 1959, we have evidence of significant breaks in 1986 and 1997. In particular, the 1997 break date may be interpreted as a signal for the retrieved monetization trend after the monetary turmoils of the previous decades, while the 1986 break date can be seen as the stabilizing effects of a macroeconomic program, the Austral plan (June 1985). The program consisted in a set of stabilization policies against the stop-ad-go cycles of argentinean high-inflation, through a reduction of fiscal deficits, a gradual de-indexation of the economy and a prices-exchange rate-wages fixing commitment.

5.3 Co-breaking

The conditional model shows an interesting case of co-breaking, *i.e.* breaks which are statistically significant in marginal models become irrelevant features when money demand equation is estimated (Hendry and Massmann, 2007; Urga, 2007). In particular, this is the case of the shocks in the occurrence of the hyperinflation episode of 1989-1990, the economic crisis of 2001-2002 or the reversion of the chronic demonetization of Argentina in act for a decade since the implementation of the Convertibility plan (1991). Evidence of co-breaking can of course be detected through the Hendry and Massmann (2007) procedure, which is confirmed in this paper by searching for breaks in money considered as a marginal process. Results, reported in Table 6, suggest that real money balances are affected by three significant deterministic breaks, in 1959, 1975 and 1991. Note that the first break is detected in conditional models but not in marginal models, while the second one is also detected in the inflation process. The third break then vanish when linear combination of the series, such as conditional models, are considered.

Insignificant estimated breaks in 1946 and 1986 appear nevertheless significant in conditional models. The effect of the macroeconomic relationship between real balances and its determinants is an evidence of the power of our empirical strategy. After a visual inspection of the money series, one would conclude that main regime changes took place in 1959 (a downward shift in money holdings), in 1975 (a change in the trend pattern) and in 1991 (the rapid re-monetization induced by the stabilization). However, results from the whole macroeconomic model suggest that the main downward shift

Table 6: Structural breaks in real M1 series

$m - p$	significant breaks				
	(i)	(ii)	(iii)	(iv)	(v)
Date	1975	1991	1959	1946	1986
PB p -value	0.05	0.01	0.03	0.30	0.21
NPB(1) p -value	0.05	0.00	0.05	0.44	0.32
NPB(2) p -value	0.05	0.00	0.05	0.40	0.34

Note: See Table 4.

in money holding can be pushed back to 1913 and the trend pattern of the money velocity changed in 1946. The break in 1991, detected when money is considered as a marginal, captures the immediate effect of the inflation stabilization and it can be seen as a level/trend adjustment after the regime of chronic monetary disorder. Breaks in 1989 and 1991 might offset each other in conditional models, probably because the one-shot peak of the hyperinflation episode in 1990 can be hardly represented through a vector of deterministic breaks. The change in monetary trends could be then assigned to other dates, like 1986 or 1997, to compensate the disappearance of the 1991 date.

6 Estimating long-run money demand equations for Argentina

6.1 Estimator and tests

We estimate long-run equations through the dynamic GLS estimator (DGLS) developed by Stock and Watson (1993). The long-run equation includes k leads and lags of first-differenced regressors. We fix $k = 2$ for all models, which is sufficient to correct for potential endogeneity in a context of annual observations. Further, each money demand model includes dummies for the breaks estimated in the previous section. Our estimator also embeds the Cochrane-Orcutt algorithm, to potentially correct for residual serial-correlation, and we set $l = 1$ the autoregressive error process. The estimated equation takes the following form:

$$(m - p)_t = \alpha + \delta t + \beta' X_t + \sum_{j=-k}^k \gamma'_j \Delta X_{t-j} + \sum_{i=1}^I (\eta_i B_{i,t}^M + \theta_i B_{i,t}^L) + \varepsilon_t \quad (7)$$

where X_t is the matrix of regressors, k is the number of leads and lags involved in the DGLS estimation and $B_{i,t}^M$ and $B_{i,t}^L$ are the vectors of deterministic breaks. Standard misspecification tests are applied to regression residuals. Under the hypothesis of stationarity of the error term, fast-double bootstrap p -values are computed through non-parametric (with d.o.f. correction) resampling (Davidson and MacKinnon, 2000,

2007)¹¹. The number of generated bootstrap samples is $B = 9,999$.

6.2 Residual-based test for cointegration

To test for cointegration, we implement the KPSS residual-based test developed by Shin (1994), which assumes the null of stationarity of residuals (cointegration) against non-stationarity. Although this test has the advantage of being designed for leads and lags regressions, it does not allow for structural breaks in its original version. Carrion-i-Silvestre and Sansò (2006) first studied the asymptotic and finite-sample performance of this test statistic when one break is included. The test statistic takes the multivariate KPSS form:

$$SC = T^{-2} \hat{\omega}^{-2} \sum_{t=1}^T \left(\sum_{j=1}^t \hat{\varepsilon}_j \right)^2 \quad (8)$$

where $\hat{\varepsilon}_t$, $\{\hat{\varepsilon}_t\}_{t=1}^T$, are the estimated DGLS residuals from Equation (7) and $\hat{\omega}^2$ is a consistent estimate of the long-run variance of $\{\varepsilon_t\}_{t=1}^T$, the latter being computed through the Bartlett kernel and a modified automatic bandwidth selection method as suggested in Kurozumi (2002)¹².

Nevertheless, the distribution of this statistic is dependent on the deterministic structure and the number of regressors involved in the model (Shin, 1994), as well as the type of break (constant, trend, stochastic parameter) and its relative position in the sample. In addition, it can also be shown that the asymptotic distribution crucially depends on the absolute number of breaks. Since our model involves up to 5 deterministic breaks in the long-run equations, we cannot use critical values provided by Carrion-i-Silvestre and Sansò (2006). In this paper, we compute instead fast-double bootstrap p -values for the multiple breaks SC statistic (MSC) through non-parametric (with d.o.f. correction) resampling.

6.3 Results

Results from the model including the interest rate as opportunity cost are reported in Table 7. We first run the unconstrained regression (1), including all breaks estimated in Section 5. Coefficients, except for the trend break in 1913, are statistically significant at

¹¹The fast-double bootstrap (FDB) is a procedure for calculating bootstrap p -values more accurately than the ordinary bootstrap method and with small computational costs. From the ordinary bootstrap p -value, $p^*(\hat{\tau}) = \frac{1}{B} \sum_{j=1}^B I(\tau_j^* > \hat{\tau})$, the FDB p -value is computed by calculating only one second-level bootstrap statistic τ_j^{**} , along with each τ_j^* , from each $j = 1, \dots, B$ first-level bootstrap samples. Let $\hat{Q}_B^{**}(1 - p^*(\hat{\tau}))$ be the $1 - p^*$ quantile of the τ_j^{**} distribution. The FDB p -value is then computed from $p^{**}(\hat{\tau}) = \frac{1}{B} \sum_{j=1}^B I(\tau_j^* > \hat{Q}_B^{**}(1 - p^*(\hat{\tau})))$.

¹²The long-run variance ($\hat{\omega}_\varepsilon^2$) is computed as sum of the estimated variance ($\hat{\gamma}_\varepsilon(0)$) plus the estimated long-run autocovariance ($2 \times \hat{\lambda}_\varepsilon$). A kernel function enters the computation of the long-run autocovariance: $\hat{\lambda}_\varepsilon = \sum_{j=1}^M w\left(\frac{j}{\tilde{M}}\right) \times \frac{1}{T} \sum_{t=j+1}^T \hat{\varepsilon}_{t-j} \hat{\varepsilon}_t$. Kurozumi (2002) proposes to estimate the kernel function $w(\cdot)$ by using a standard Bartlett window, but imposing a modified Andrews (1991) automatic lag length (\tilde{M}) selection method: $\tilde{M} = \min\left(1.1447 \left\{ \frac{4\hat{\rho}^2 T}{(1+\hat{\rho})^2(1-\hat{\rho})^2} \right\}^{1/3}, 1.1447 \left\{ \frac{4k^2 T}{(1+k)^2(1-k)^2} \right\}^{1/3}\right)$, where $\hat{\rho}$ is the AR(1) coefficient of $\hat{\varepsilon}_t$ and k can take a value between 0.7 and 0.9. In this paper we select $k = 0.8$, as suggested by Carrion-i-Silvestre and Sansò (2006).

1, 5 and 10%. In particular, income and interest rate parameters are both highly significant and with expected signs. The former is positive and close to unity, in contrast with the preliminary cointegration results reported in Table 2, while the latter is around -1.6. These results suggest that the long-run demand for real money balances in Argentina is driven by proportional changes in income (unit homogeneity hypothesis), while portfolio motives lead to strong adjustments of the desired level of money to variations in the yield of alternative assets. It is worth noticing that results are in line with those reported for industrialized countries (see, among others, Lucas, 1988, and Hendry and Ericsson, 1991). However, the inclusion of structural breaks is here a crucial long-run issue. Estimated residuals have good statistical properties, except for some evidence of ARCH effects at lag 2 and 4. In addition, the residual-based *MSC* test cannot reject the null hypothesis of cointegration.

In regression (2), we constrain the insignificant break coefficient to zero and we estimate the model. Results are very similar to those obtained in regression (1). We then test for the income homogeneity hypothesis through a Wald test. The null hypothesis being not rejected (p -value = 0.77), we estimate regression (3) including this restriction over the real GDP elasticity. Results are very similar to those reported in regressions (1) and (2). As expected, the level break in 1913 is negative (the end of the gold-standard system), while the level break in 1946 is positive (a transitory effect of the redistributive peronist policies). The level break in 1959 has a negative sign, which can be associated to the fall of real balances due to sudden inflationary shocks which took place in that year. It is worth noticing that the coefficient for the trend break in 1946 has a negative sign, which identifies the beginning of the long-lasting demonetization of Argentina until the 1990s. The positive sign of the coefficient for the trend break in 1959 partially compensate the fall of money since the half of the 20th century (the sum of the coefficient is still negative). In regression (4) and (5) we allow for a second lag in the Cochrane-Orcutt procedure to check the robustness of previous estimates. As expected, results are quantitatively similar to those in regression (2) and (3). Again, residual analysis does not highlight major problems in our specifications, except for some evidence of ARCH effects at lag 4, and cointegration cannot be rejected. Error-correction terms from regressions (3) and (5), *i.e.* the estimated deviations of real money balances from the long-run equilibrium, are plotted in Figure 3(a).

Results from the model including the inflation rate as opportunity cost are reported in Table 8. We first run the unconstrained regression (1). Income and inflation coefficients are highly significant and with expected signs, while many level and trend break parameters appear insignificant. Unlike the previous specification, here the income coefficient (1.48) is significantly higher than unity and the inflation coefficient (-0.59) is quite lower than the interest rate coefficient. The latter issue confirms results showed in Table 2 about the existence of a weak Fisher parity between interest rate and inflation. Estimated residual have good statistical properties, except for some autocorrelation at lag 2, and the *MSC* test cannot reject the null hypothesis of cointegration. In regression (2) and (3), we look for a more parsimonious specification by constraining insignificant breaks to zero. It is worth noticing that estimated income and inflation coefficients slightly differ from the baseline regression (1): the former tends to be lower, while the latter stays in the range of the confidence interval. We test for the income homogeneity hypothesis. Although the high income coefficient, the null hypothesis cannot be

Table 7: Long-run equation for the model $(m - p, y, r)$

	(1)	(2)	(3)	(4)	(5)
y	0.92***	0.94***	1	1.03***	1
r	-1.59***	-1.59***	-1.61***	-1.62***	-1.62***
1913 ^M	-0.23**	-0.21***	-0.21***	-0.20***	-0.20***
1913 ^L	-0.01	-	-	-	-
1946 ^M	0.32***	0.31***	0.31***	0.35***	0.35***
1946 ^L	-0.03***	-0.03***	-0.03***	-0.03***	-0.03***
1959 ^M	-0.37***	-0.37***	-0.38***	-0.39***	-0.38***
1959 ^L	0.02*	0.02*	0.02**	0.02**	0.02**
Constant	-0.05	-0.32	-1.30	-1.85	-1.29
Trend	0.009	-0.001	-0.003	-0.005	-0.004
Obs.	102	102	102	102	102
k	2	2	2	2	2
l	1	1	1	2	2
χ^2 Restrictions	-	-	0.084	-	0.045
			[0.77]	-	[0.83]
MSC	0.0236	0.0242	0.0247	0.0277	0.0279
	[0.36]	[0.39]	[0.36]	[0.13]	[0.12]
AR (1)	[0.41]	[0.38]	[0.37]	[0.93]	[0.85]
AR (2)	[0.11]	[0.11]	[0.10]	[0.99]	[0.99]
AR (4)	[0.21]	[0.19]	[0.18]	[0.98]	[0.98]
J-B	[0.61]	[0.61]	[0.62]	[0.99]	[0.99]
B-P	[0.24]	[0.22]	[0.21]	[0.11]	[0.12]
ARCH (1)	[0.10]	[0.10]	[0.11]	[0.38]	[0.35]
ARCH (2)	[0.01]	[0.01]	[0.01]	[0.10]	[0.11]
ARCH (4)	[0.01]	[0.01]	[0.01]	[0.02]	[0.02]

Note: DGLS estimations. Upper-scripts M and L are break in the mean and in the linear trend, respectively. $AR(p)$ is the Breusch-Godfrey statistic, B-P is the Breusch-Pagan statistic, J-B is the Jarque-Bera statistic, $ARCH(p)$ is the Engle-Bollerslev statistic, MSC is the residual-based test for cointegration with multiple breaks. Standard errors are reported in parentheses. Fast-double bootstrap p -values in square brackets, simulated through a non-parametric sampling with 9,999 random replications. Asymptotic p -values for linear restrictions hypothesis testing in square brackets.

***, ** and * denote significance at 1, 5 and 10%, respectively.

rejected in both specifications (2) and (3) (p -value = 0.43 and 0.28, respectively). We then estimate constrained regressions (4) and (5). Results are similar to those reported in regressions (2) and (3). In particular, we estimate the inflation coefficient in a range between -0.6 and -0.7. Residual tests do not point out any statistical problem and cointegration hypothesis is never rejected. Error-correction terms from regressions (4) and (5) are plotted in Figure 3(b).

As before, the level break in 1913 is negative and the level break in 1946 is positive. Again, the level break in 1959 has a negative sign, as well as the trend break in 1946. However, the trend break in 1959 is not significant in all specifications.

Table 8: Long-run equation for the model ($m - p, y, \pi$)

	(1)	(2)	(3)	(4)	(5)
y	1.48***	1.23***	1.32***	1	1
π	-0.59***	-0.69***	-0.55***	-0.70***	-0.59***
1913 ^M	-0.21**	-0.23***	-0.22***	-0.24***	-0.23***
1913 ^L	0.006	-	-	-	-
1946 ^M	0.27***	0.29***	0.33***	0.29***	0.33***
1946 ^L	-0.02**	-0.03***	-0.04***	-0.03***	-0.03***
1959 ^M	-0.30***	-0.31***	-0.26***	-0.29***	-0.26***
1959 ^L	-0.01	-	-	-	-
1986 ^M	0.20*	0.25***	-	0.21**	-
1986 ^L	0.04	-	0.06***	-	0.04***
1997 ^M	-0.04	-	-	-	-
1997 ^L	0.04	0.07	-	0.06***	-
Constant	-9.25*	-5.15	-6.66	-1.44	-1.43
Trend	-0.021	-0.007	-0.010	0.001	0.001
Obs.	102	102	102	102	102
k	2	2	2	2	2
l	1	1	1	1	1
χ^2 Restrictions	-	-	-	0.62	1.16
				[0.43]	[0.28]
MSC	0.0313	0.0345	0.0361	0.0327	0.0345
	[0.07]	[0.18]	[0.13]	[0.22]	[0.15]
AR (1)	[0.24]	[0.17]	[0.15]	[0.11]	[0.07]
AR (2)	[0.04]	[0.03]	[0.02]	[0.02]	[0.01]
AR (4)	[0.17]	[0.14]	[0.08]	[0.12]	[0.06]
J-B	[0.46]	[0.45]	[0.47]	[0.45]	[0.47]
B-P	[0.24]	[0.31]	[0.06]	[0.28]	[0.05]
ARCH (1)	[0.32]	[0.25]	[0.25]	[0.15]	[0.09]
ARCH (2)	[0.16]	[0.12]	[0.22]	[0.17]	[0.19]
ARCH (4)	[0.30]	[0.26]	[0.51]	[0.31]	[0.42]

Note: See Table 7.

[Figure 3 about here]

7 Short-run dynamics

To model dynamics, we follow Engle and Granger (1987) and we estimate the following general autoregressive specification for the growth rate of money:

$$\Delta(m - p)_t = \alpha_0 + A(L)\Delta(m - p)_t + C(L)\Delta\mathbf{x}_t + \gamma ECT_{t-1} + \phi' \Delta\mathbf{B}^{M,L} + u_t \quad (9)$$

where $A(L)$ is a polynomial vector and $\mathbf{B}(L)$ is a polynomial matrix, $\Delta\mathbf{x} = \{\Delta y, \Delta r, \Delta\pi\}'$

is the vector of regressors, $\Delta B^{M,L}$ is the vector of structural breaks¹³ and ECT is the error-correction term.

We estimate Equation (9) using up to 4 lags. A parsimonious single-equation error-correction model is then obtained by dropping insignificant variables and following the automated General-to-Specific reduction approach (Hendry and Krolzig, 2001; Doornik, 2007). In addition, we report standard diagnostic tests and we allow for outliers correction through the introduction of blip-dummy variables for large residuals. Asymptotic significance at 1, 5 and 10% (denoted by ***, ** and * , respectively) is reported, while t -statistics for coefficients significance are based on heteroscedasticity and autocorrelation consistent standard errors (Andrews, 1991). As for the long-run specification, two models involving either the interest rate or the inflation rate are estimated. Error-correction terms from specification (3) in Table 7 and from specifications (4) and (5) in Table 8 are used.

Following this strategy, we report in Equation (10) our preferred dynamic equation for the model involving the growth rate of GDP and changes in the interest rate:

$$\begin{aligned} \Delta(m-p)_t = & 0.11^* \Delta(m-p)_{t-1} - 0.16^{***} \Delta(m-p)_{t-2} + 0.79^{***} \Delta y_t \\ & - 0.84^{***} \Delta r_t - 0.47^{***} ECT_{t-1} - 0.21^{***} \Delta 1913^M \\ & + 0.13^{***} \Delta 1946^M - 0.29^{***} \Delta 1959^M \end{aligned} \quad (10)$$

$$\begin{array}{ccccccc} T=102 & R^2=0.76 & \hat{\sigma}=9\% & SBC=-1.69 & & & \\ AR(4)=1.69 & ARCH(4)=0.79 & NORM=0.66 & HET=1.09 & & & \end{array}$$

where $AR(p)$ is the Breusch-Godfrey LM test for residual serial-correlation, $ARCH(p)$ is the Engle-Bollerslev test for the presence of ARCH effects, NORM is the Doornik-Hansen test for normality of residuals and HET is the White test for heteroscedasticity. Although only a small set of regressors enters the final dynamic model, the equation passes all the diagnostic tests. Our preferred model includes 3 (differenced) level breaks (1913, 1946 and 1959) and one outlier for the year 1989.

We finally report our preferred dynamic equations for the model involving the growth rate of GDP and the acceleration in inflation. Equation (11) and (12) use specifications (4) and (5) from Table 8, respectively:

$$\begin{aligned} \Delta(m-p)_t = & 0.17^{***} \Delta(m-p)_{t-1} + 0.84^{***} \Delta y_t - 0.35^{***} \Delta \pi_t - 0.28^{***} ECT_{t-1} \\ & - 0.25^{***} \Delta 1913^M + 0.13^{***} \Delta 1946^M - 0.11^{***} \Delta 1959^M \\ & - 0.03^{**} \Delta 1946^L + 0.06^{***} \Delta 1997^L \end{aligned} \quad (11)$$

$$\begin{array}{ccccccc} T=102 & R^2=0.80 & \hat{\sigma}=8\% & SBC=-1.82 & & & \\ AR(4)=2.27^* & ARCH(4)=2.41^* & NORM=0.62 & HET=1.02 & & & \end{array}$$

¹³Breaks enter the short-run model in first-difference because they were included into the long-run equation.

$$\Delta(m-p)_t = 0.18^{***}\Delta(m-p)_{t-1} + 0.71^{***}\Delta y_t - 0.34^{***}\Delta\pi_t - 0.35^{***}ECT_{t-1} - 0.25^{***}\Delta 1913^M + 0.12^{***}\Delta 1946^M - 0.17^{***}\Delta 1959^M \quad (12)$$

$$T=102 \quad R^2=0.81 \quad \hat{\sigma}=8\% \quad SBC=-1.91 \\ AR(4)=1.58 \quad ARCH(4)=0.93 \quad NORM=2.19 \quad HET=1.61$$

Only one lag for the money growth term enters the dynamic equation. Level breaks in 1913, 1946 and 1959 are again significant, together with two additional trend breaks (1946 and 1997) for the specification (4). For the latter, we have some evidence of residual autocorrelation and conditional heteroscedasticity, although only at 10% level of significance. One outlier for the year 1989 enters both specifications, while an additional outlier for the year 1974 enters Equation (12).

Turning to the dynamic adjustment of real money balances, the error-correction term is significant in all preferred equations and with the expected (negative) sign. It is worth noticing that the speed of adjustment towards the long-run equilibrium is stronger in Equation (10). This is in line with results obtained in Section 6, where the estimated long-run interest rate coefficient is significantly larger than the long-run inflation coefficient. The overall high explained variance, which is close to 80%, suggests that our simple models are good enough to fit actual real money data. To further explore this issue, we plot fitted and actual values of $\Delta(m-p)_t$ from Equations (10), (11) and (12) in Figure 4(a), 4(b) and 4(c), respectively.

[Figure 6 about here]

8 Concluding Remarks

In this paper, we estimated money demand in Argentina over a long time span, from 1900 to 2004. When we tested for cointegration between real M1 balances, real GDP, inflation rate and interest rate, we showed that cointegration results are not robust when structural breaks are neglected. We then tested and dated for deterministic structural breaks in marginal (real GDP, inflation rate and interest rate) and conditional (real M1 balances) processes using a sequential search procedure. Results suggested the presence of significant breaks both in marginal and conditional processes, consistent with extreme financial episodes and macroeconomic shocks. These breaks coincides with the end of the gold-standard experience of Argentina (1913), the beginning of Peronist policies (1946), the explosion of fiscal deficits and prices (1959), the attempts of stabilization as a consequence of the Austral Plan (1986) and the re-monetization process during the 90s (1997).

We estimated long-run money demand equations embedding structural breaks by DGLS. Estimated parameters are highly significant and correctly signed. Quite interestingly, the quantitative theory hypothesis of income homogeneity could not be rejected. Conclusive evidence on the existence of long-run money demand equilibrium in Argentina is provided by multiple breaks cointegration tests. In addition, although the presence of extreme adjustments which took place in 1975 and in 1989-1990, estimated cointegrating vectors appear generally stable once the breaks are taken into account. Finally, the short-run analysis highlighted the high speed at which agents re-adjusted

excess money holdings towards the desired level of liquidity. These findings are in line with those reported in the literature of industrialized countries (Lucas, 1988; Hendry and Ericsson, 1991; Muscatelli and Spinelli, 2000), but add new elements to the understanding of the long term monetary dynamic in emerging economies.

Some interesting issues arise. The welfare cost of inflation (Lucas, 2000) and the seignorage cost of protracted monetary disorder can be straightforwardly explored by simply exploiting the results reported in this paper. In addition, the moderate re-monetization process in Argentina after decades of financial chronic disorder, as suggested by data and by a quite slower speed of adjustment in the post-1991 period, may also point out the effects of monetary hysteresis phenomena which need to be further explored. Indeed, results reported in this paper call for a comparative investigation extended to other Latin American countries. We leave these issues for future research.

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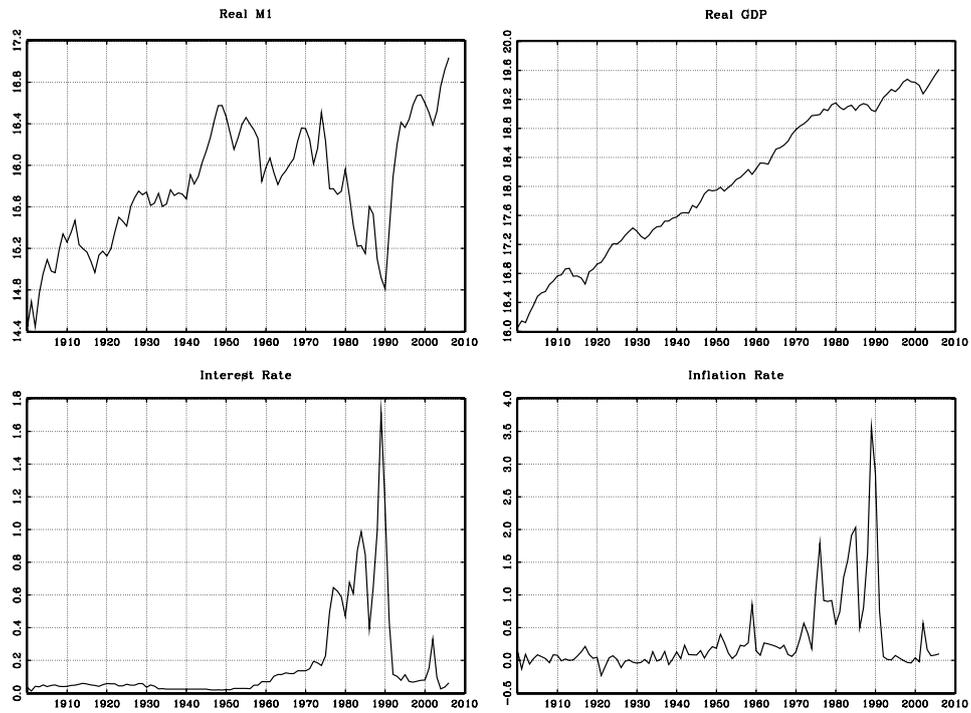
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9 Data Appendix

Many sources were used to build the dataset. The main one is Veganzones and Winograd (1996), who collected data until 1992. From 1993 to 2006, the dataset is fully updated by the authors of the present work, while time series before 1993 are corrected and re-scaled.

- **Money** (*Nominal M1*) :
1900-1913 : Della Paolera (1988) and Anuario Geografico Argentino;
1913-1935 : Anuario Geografico Argentino (various issues);
1935-2006 : Banco Central de la Republica Argentina (*memorias* and web sources).
- **Gross Domestic Product** (*Nominal*) :
1900-1993 : Della Paolera and Taylor (2003);
1993-2006 : International Monetary Fund (International Financial Statistics).
- **Interest Rate** (*Various Rates*) :
1900-1913 : Della Paolera (1988);
1913-1938 : Anuario Geografico Argentino;
1938-2006 : Banco Central de la Republica Argentina (*memorias* and web sources).
- **Price Index** (*Wholesale Price Index, base-year 1993*) :
1899-1956 : Della Paolera (1988), Economic Commission for Latin America (ECLA), United Nations and Instituto Nacional de Estadistica y Censos (INDEC).
1956-2006 : Banco Central de la Republica Argentina (web sources).
- **Implicit GDP Price Deflator** (*Base-year 1993*) :
1900-1993 : Della Paolera and Taylor (2003);
1993-2006 : International Monetary Fund (International Financial Statistics).
- **Gross Domestic Product** (*Real*) :
1900-1980 : Because of rebasements in 1970 and 1986, an adjustment was needed to harmonize the real GDP series. This was done by linking the pre-1980 series to the post-1980 series. We applied growth rates of the real GDP, originally computed as the nominal GDP deflated by the implicit price index, to the 1980 value (issued from Della Paolera and Taylor, 2003) and we adjusted the series backward;
1980-1993 : Della Paolera and Taylor (2003);
1993-2006 : International Monetary Fund (International Financial Statistics).

Figure 1: Argentinean data (1900-2006)



Notes: Money balances and GDP are expressed in Argentinean pesos.
Sources: see Appendix

Figure 2: Log of money velocity ($v = y - m + p$)

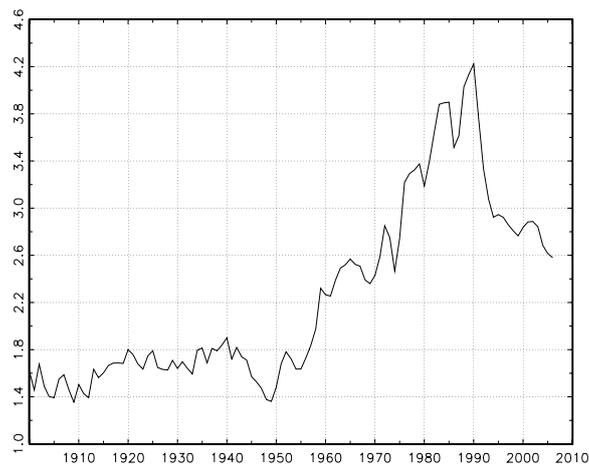
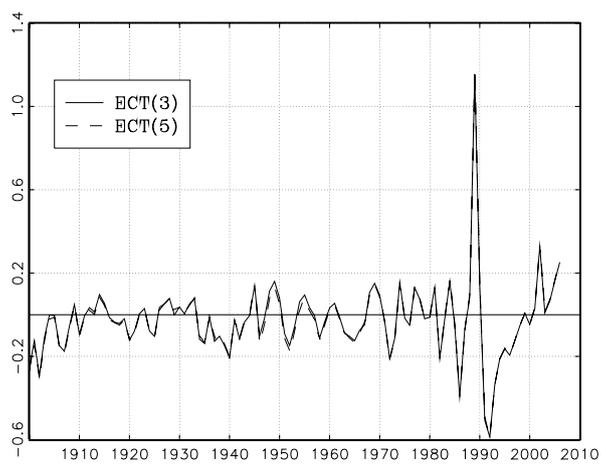
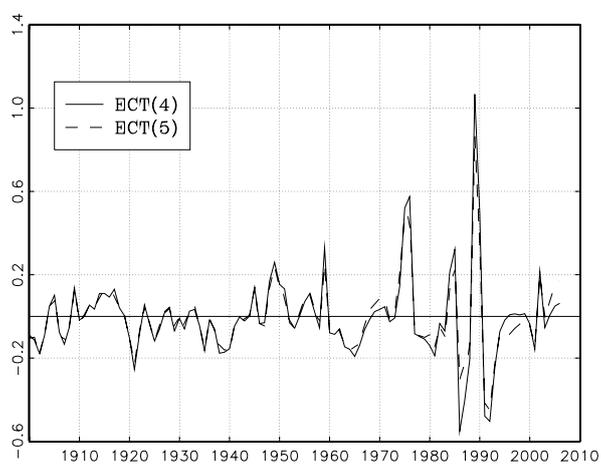


Figure 3: Deviations from the long-run equilibrium for real balances.

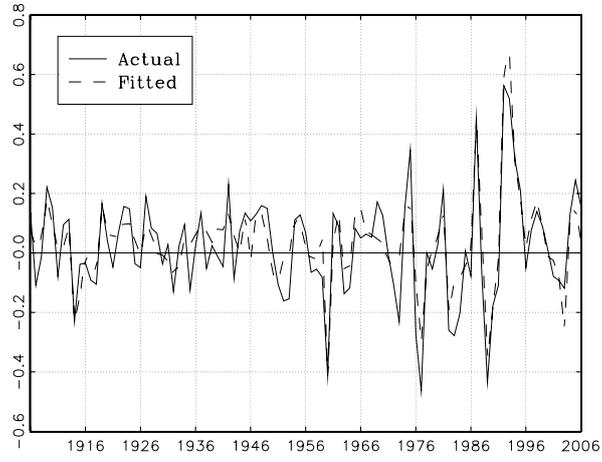


(a) Error-correction term (ECT) (3) and (5) from Table 7.

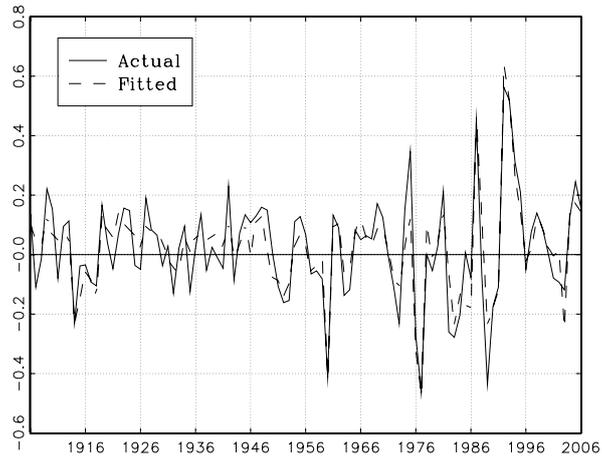


(b) Error-correction term (ECT) (4) and (5) from Table 8.

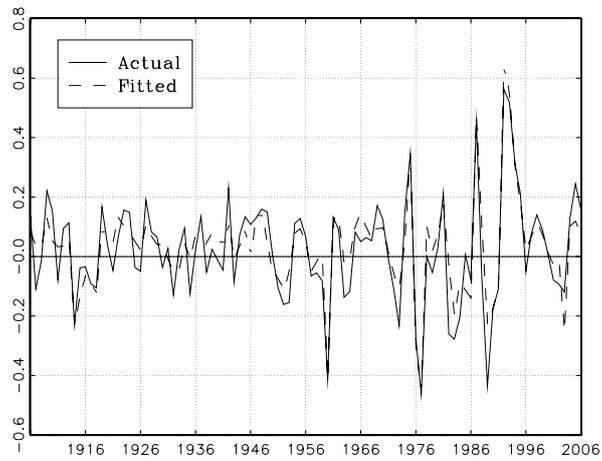
Figure 4: Actual and Fitted Values for $\Delta(m - p)_t$



(a) Equation (10).



(b) Equation (11).



(c) Equation (12).