European Economic Review 187 (2026) 105359

Contents lists available at ScienceDirect

European Economic Review

journal homepage: www.elsevier.com/locate/eer

Inflation targeting, Monetary policy, and Inequality

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ARTICLE INFO

JEL classification:
D31
E4
E5

Keywords:
Inflation targeting
Income inequality
Monetary policy

ABSTRACT

This paper explores whether the transmission from monetary policy to income inequality may depend on the adoption of Inflation Targeting (IT) regimes. Using an interacted panel VAR, we find that expansionary monetary policy shocks reduce income inequality in countries that have switched to IT regimes. In contrast, in non-IT regimes the same shock is associated with a short-lived increase in income inequality. A decomposition of transmission channels indicates that the employment channel is the primary equalizing mechanism under IT, as expansionary shocks generate stronger improvements in labor market conditions. The financial channel operates in the opposite direction but is quantitatively smaller. We further show that the inequality-reducing effects of monetary policy are not replicated by other institutional features often associated with credibility, such as central bank transparency or central bank independence. Our findings are robust to alternative identification schemes, broader classifications of IT regimes, controls for self-selection into IT adoption, and to conditioning on different inflation environments.

1. Introduction

Choosing a monetary policy regime is a perennial issue for policymakers, regardless of the challenges faced by monetary policy. Inflation Targeting (IT) has emerged as a polular monetary policy regime, adopted by more than 20 central banks (CB) around the world.[1] Since its adoption in the early 1990s, concerns have been expressed about the distributional and employment consequences of IT. Akerlof et al. (1996) argue that targeting very low inflation can result in a sustainable but unnecessarily high rate of unemployment, thus exacerbating income inequality.[2] Epstein (2007) points out that IT creates a culture of "inflation obsession", leading CBs to neglect other social objectives, while Stiglitz (2012) criticizes IT frameworks for potentially exacerbating inequality by neglecting financial stability objectives and allowing high levels of unemployment.

Despite the concerns regarding the social impact of IT, the existing empirical literature primarily focuses on its effects on inflation and inflation expectations, with limited evidence on its impact on employment or the distribution of income. To the best of our knowledge, no study has examined how monetary frameworks, such as IT, influence the transmission of monetary policy shocks

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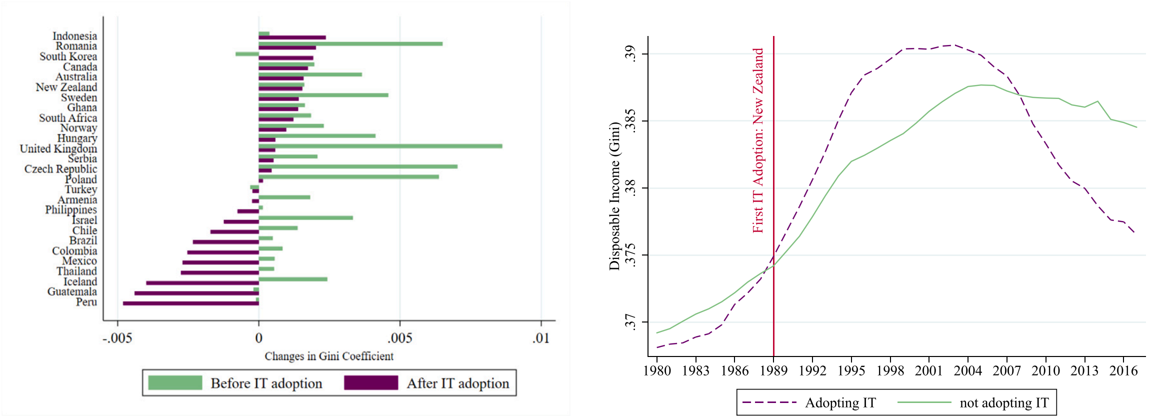
[1] IT has been implemented in various forms, with a balance between discretion, constraints, and rules. This balance ranges from discretionary or "traditional" regimes to targeting regimes based on money targets, price targets, nominal income targets, exchange rates, and fixed exchange rates (Ball, 2010).

[2] The authors simulate a long-run Phillips curve and demonstrate significant increases in unemployment and decreases in real wages when authorities target inflation rates below 3%.

https://doi.org/10.1016/j.euroecorev.2026.105359
Received 13 January 2025; Received in revised form 25 April 2026; Accepted 30 April 2026

Available online 5 May 2026
0014-2921/



(a) Inequality changes before and after the adoption of Inflation Targeting (1980–2017).

(b) Average inequality levels in IT and non-IT countries.

Fig. 1. Income inequality dynamics and inflation Targeting regimes.
Note: Income inequality is measured using the Gini coefficient of disposable income from the Standardized World Income Inequality Database (SWIID). Countries classified as IT adopters follow the definition in Hammond (2012).

to inequality. This paper aims to fill this gap. Robust evidence on the potential role of IT in shaping inequality through monetary policy can help safeguard central banks from political risks, such as populist demands to adopt quasi-fiscal measures to mitigate inequality, which fall outside their mandate and undermine their credibility (Rogoff, 2021).

The distributional implications of IT have attracted increasing attention from policymakers and economists. In a recent speech, da Silva (2020) notes that income inequality has risen in both IT and non-IT countries. However, focusing on countries that transitioned from non-IT to IT regimes, as shown in Fig. 1(a), reveals sizable reductions in inequality changes following adoption in several cases. Moreover, in countries where inequality increased both before and after IT adoption, these increases are markedly smaller after the transition. Fig. 1(b) compares inequality trends in 27 IT adopters with those in 49 economies that never adopted IT. Until the late 1980s, both groups exhibited similar inequality levels, but a divergence emerges as countries begin adopting IT, starting with New Zealand in 1989. Taken together, this exploratory evidence suggests that the distributional consequences of switching to IT regimes warrant further investigation.

The relationship between IT and inequality can be influenced by third factors which include an economy's fiscal stance, trade and financial interdependence, exchange rate regime, as well as institutional characteristics, among others. These factors can impact both the decision to adopt an IT regime and inequality trends. Since IT is predominantly concerned with the conduct of monetary policy, the key question to explore is whether the presence of IT influences the transmission of monetary policy to income distribution over and above the role of additional confounding factors.

To this aim, we draw upon the expanding literature on the distributional consequences of monetary policy (Coibion et al., 2017; Mumtaz and Theophilopoulou, 2017; Evgenidis and Fasianos, 2021). The evidence from this literature remains inconclusive. While several studies find equalizing effects of monetary policy shocks, others find no statistically significant association, or even contradicting results depending on the country and period under investigation, or the nature of the monetary policy shock (Colciago et al., 2019). Although numerous studies have assessed the impact of IT on macroeconomic outcomes such as inflation and inflation volatility (Mishkin and Schmidt-Hebbel, 2007; Lin and Ye, 2007; Gambetti and Pappa, 2009; Lin and Ye, 2009), output and output fluctuations (Ball and Sheridan, 2004; Gonçalves and Carvalho, 2009), and exchange rates (López-Villavicencio and Pourroy, 2019), the role of IT in shaping distributional outcomes has not been explored.

The primary contribution of our paper is to bridge these two distinct strands of literature: the strand exploring the macroeconomic consequences of IT and the more recent one on the distributional impacts of monetary policy, with the aim of examining how the impact of monetary policy on income inequality is influenced by the adoption of IT regimes. To achieve this, we utilize a Panel VAR model for a sample of 76 developed and developing economies, 27 of which have transitioned to an IT regime. Therefore, our paper relies on a large panel sample, comprising both emerging and developed economies, which is relatively scarce in the literature exploring the distributional impact of monetary policy.[3]

Modeling how IT shapes macroeconomic dynamics using panel VAR models has received limited attention. An exception is Mishkin and Schmidt-Hebbel (2007), who employ a conventional panel VAR to examine whether IT helps countries achieve lower inflation in the long run. However, such approaches estimate separate models for IT and non-IT countries, implicitly treating

[3] Furceri et al. (2018) is one of the few studies that perform such an analysis with both developed and developing economies, though not focusing on the monetary policy framework employed in each country.

regime status as fixed and overlooking the fact that IT adopters transition from non-IT regimes. We address this limitation by estimating an interacted panel VAR (IP-VAR) in which model parameters differ systematically across IT and non-IT regimes within countries over time. This unified framework allows a direct comparison of impulse-response functions (IRFs) across regimes and enables us to assess how the dynamic relationship between monetary policy and income inequality varies with the adoption of IT. While interaction terms are common in single-equation models, their application in VARs was introduced by Loayza and Raddatz (2007), Towbin and Weber (2013), and Sá et al. (2014).[4] In contrast to this literature, we focus on IT as a monetary policy framework and its role in the transmission of expansionary monetary policy shocks to income inequality.

Our baseline specification is a medium-scale interacted panel VAR that incorporates a rich macroeconomic and financial information set. Moving beyond small-scale VARs, the model includes a wider range of variables, allowing us to capture how monetary policy shocks propagate through the economy and translate into distributional outcomes under different IT regimes. The use of a medium-scale VAR serves two main purposes. First, it allows us to go beyond a narrow core set of variables, often limited to GDP and inflation, and to capture additional macroeconomic and financial effects associated with monetary policy, including fiscal outcomes, labor market conditions, credit spreads, asset prices, and housing market dynamics. Second, conditioning on a richer information set strengthens the identification of monetary policy shocks, thereby mitigating the limitations of recursive identification in low-frequency data.

Beyond aggregate macroeconomic dynamics, our paper also contributes to the literature by examining how IT affects the transmission of monetary policy through distributional channels. Two main channels emerge from the literature, both linked to the income composition of heterogeneous households: the employment channel and the financial channel. The employment channel reflects the greater reliance of lower-income households on labor income (Coibion et al., 2017; Furceri et al., 2018; Dolado et al., 2021; Samarina and Nguyen, 2024). Expansionary monetary policy that stimulates employment and wage growth disproportionately benefits these households, leading to a reduction in income inequality. Conversely, contractionary policy reduces labor demand, prolongs unemployment, and slows wage growth, adversely affecting households at the lower end of the income distribution (Gornemann et al., 2021). The financial channel operates in the opposite direction. Expansionary monetary policy raises asset prices and capital gains, disproportionately benefiting wealthier households that derive income from asset holdings (O'Farrell and Rawdanowicz, 2017; Leroy and Lucotte, 2019). Moreover, if expansionary policy generates inflationary pressures, it may erode the real incomes of households holding non-indexed assets, further exacerbating income inequality (Doepke and Schneider, 2006).

Our hypothesis is that the presence of IT influences these distributional channels of monetary policy. We argue that, following expansionary monetary policy shocks, an IT framework: (i) strengthens the employment channel of monetary policy transmission and (ii) mitigates the erosion of non-indexed real incomes by effectively controlling inflation. These aspects have not been previously explored, distinguishing our study from the existing literature.

A key challenge in empirically assessing these hypotheses is that the adoption of an IT framework is not random. Countries may self-select into inflation targeting based on characteristics that also influence income inequality, potentially biasing the estimated effects. To address this concern, our analysis explicitly accounts for self-selection bias. Specifically, this paper is the first in the relevant literature to employ propensity score matching (PSM) to assess whether the distributional effects attributed to IT adoption are driven by the regime itself rather than by confounding factors. While existing studies use PSM to examine whether differences in macroeconomic outcomes, such as inflation, exchange rate pass-through, and oil shock pass-through, reflect IT adoption or underlying characteristics (Lin and Ye, 2007, 2009; López-Villavicencio and Pourroy, 2022), none investigates the effects of IT on income inequality.

The main finding of our paper is that expansionary monetary policy shocks significantly reduce income inequality in countries operating under IT regimes. Importantly, this result does not extend to alternative institutional arrangements commonly associated with monetary policy credibility, such as higher central bank transparency or greater central bank independence, which do not generate comparable inequality reducing effects. By contrast, the response of income inequality to the same monetary policy shock in non-IT countries is economically small and largely statistically insignificant over the forecast horizon. Although inequality displays a brief positive response immediately following the shock, this effect is short-lived and quickly becomes indistinguishable from zero. Our PSM results are consistent with this finding. After correcting for potential self-selection, expansionary monetary policy shocks continue to reduce income inequality under IT, reinforcing the interpretation that the observed effects reflect the regime itself rather than underlying country characteristics. Finally, we provide evidence on the mechanisms underlying these results. We find that the employment channel is the primary equalizing force: under IT regimes, the reduction in income inequality following expansionary monetary policy shocks becomes more pronounced, as central banks' focus on price stability leads to additional employment and output gains. The financial channel is found to dampen this effect, but its quantitative impact is smaller and statistically insignificant.

The remaining sections of the article are organized as follows: Section 2 discusses the data and explains the empirical strategy. Section 3 presents the results on income inequality under different monetary policy frameworks and controls for different institutional characteristics. This section also demonstrates the robustness of our main results through alternative specifications. Section 4 addresses potential self-selection into IT using PSM methods. Section 5 concludes.

[4] These studies examine how structural characteristics shape the transmission of trade shocks, external shocks under different exchange rate regimes, and capital inflow shocks. Interaction-based VARs have since been applied to analyze the effects of fiscal shocks (Huidrom et al., 2020), monetary policy shocks (Hjortsoe et al., 2018), credit shocks (Leroy and Lucotte, 2019), and oil price and global demand shocks (Abbritti and Weber, 2018; Dkabrowski et al., 2022).

2. Data and empirical methodology

2.1. Data sources and variable description

Our sample is an annual panel setting ranging from 1980 to 2017. It comprises 27 countries that have switched to an IT regime and 49 non-IT developed and emerging economies as a control group.

We identify transitions to an IT regime drawing on Hammond (2012), Rose (2007) and Huang et al. (2019) who define IT as a monetary policy framework with the following characteristics: (i) price stability is explicitly recognized as the main goal of monetary policy; (ii) there is a public announcement of a quantitative target for inflation; (iii) monetary policy is based on a wide set of information, including an inflation forecast; (iv) transparency and accountability mechanisms are in place. To account for the transition, we assign the value of 1 to the year that countries switched to an IT regime as defined above. Note that this definition allows us to identify the time that economies de facto adopted an IT regime, which does not necessarily coincide with the year of their official announcement.[5]

To achieve a reasonable comparison between IT and non-IT countries, we have carefully selected the non-IT group comprising both developed and emerging economies. The countries included in this group have a real GDP per capita over the entire period that is equal to or greater than the poorest IT country and a population size that is equal to or larger than the smallest IT country in our non-IT group.[6] Table A.1. in the online appendix provides a list of the 27 IT countries and the 49 non-IT countries that meet these criteria, along with descriptive statistics for our inequality measures and the key macroeconomic variables employed in our analysis. It also indicates the IT adoption year, if applicable.

The following paragraphs describe the sources and transformations of the variables used in our empirical model. Our main source of data on income inequality across countries is the Standardized World Income Inequality Database (SWIID) (Solt, 2020), which is the most comprehensive cross-country dataset on income inequality.[7] We primarily rely on annual data for estimates of Gini coefficients for equivalised household net income. While the original SWIID dataset contains information on countries for a period of over 100 years for a few developed economies, we limit our sample to the period starting in 1980. This allows us to maximize the sample of countries used, including both developed and developing economies, while capturing the transition to an IT regime that began in the early 1990s.

Our proxy for monetary policy is the short-term nominal interest rate. Short-term interest rates are typically closely related to the official policy rate controlled by central banks (CBs) and are commonly used in empirical studies employing panel models of advanced and emerging economies (Elbourne and de Haan, 2006; Eleftheriou et al., 2006; Bhattacharya and Jain, 2020; Floro and Van Roye, 2017). We source the interest-rate data from the Organisation for Economic Co-operation and Development (OECD) and the official central bank websites of the respective countries for all advanced economies in our sample. For some emerging economies for which short-term rates were not available, we used the lending rate sourced from the International Financial Statistics database.[8] Furthermore, to account for the period when monetary policy reached the zero lower bound (ZLB), we used the short-term shadow rate (SSR) from Krippner (2020).[9] The shadow rate allows us to address the discontinuity in the dataset introduced by the ZLB when central banks implemented unconventional monetary policy measures.

In terms of the macroeconomic variables, we begin with a core set required to identify monetary policy shocks. Given that the primary objective of monetary policy is price stability, we follow (Mishkin and Schmidt-Hebbel, 2007) and include a measure of the inflation gap, constructed as the deviation of CPI inflation from its target for inflation-targeting countries and from a Hodrick–Prescott-filtered trend for non-targeters.[10][11] We also include real GDP to capture the effects of monetary policy on economic

[5] Consistent with the criteria outlined above, we exclude both the United States and euro area members from the group of IT countries. Although the Federal Reserve adopted an explicit inflation objective in 2012, its policy framework does not meet condition (i) as defined by Hammond (2012), namely, that price stability must be the sole objective of monetary policy. Instead, the Fed operates under a dual mandate, placing equal emphasis on maximum employment and stable prices.

As for the euro area, member states are part of a currency union and are thus classified as operating under a hard peg with no separate legal tender (Ilzetzki et al., 2022). Furthermore, stricter IT classifications require IT central banks to use the policy rate as the primary monetary policy instrument, with no concurrent target for monetary aggregates or the exchange rate (Duncan et al., 2022). The ECB's operational strategy during its first two decades has been based, at least officially, on a two-pillar framework that combines both an economic analysis and a monetary analysis, with the latter giving weight to the growth of monetary aggregates such as M3 (Gros and Capolongo, 2019).

The most significant revision to the ECB's framework came with the 2021 Strategy Review, which introduced a symmetric inflation target of 2% over the medium term. For this reason, euro area countries are also classified as non-IT countries by Hammond (2012), as well as in more recent IT classifications (Duncan et al., 2022; Ilzetzki et al., 2022; Zhang and Wang, 2022; Zhang, 2025). In the online Appendix, we incorporate additional countries under broader IT classifications and the evidence produced confirm the robustness of our main findings.

[6] A similar sampling strategy is employed by Lin and Ye (2009).

[7] The dataset has been extensively used by researchers aiming to track inequality for developed and developing countries in a comparable manner across time. See, for example, Heathcote et al. (2017), Gozgor and Ranjan (2017) and Berg et al. (2018), among others.

[8] As our sample includes many developing countries in earlier periods, obtaining the official policy rates is not always feasible. Following Floro and Van Roye (2017), we choose the lending rate as a proxy for the interest rate set by the central bank for data completeness purposes. Indeed, for the periods where we could compare the available lending rates with the policy rates or the shadow rate during zero lower bound period, we found a strong correlation of 0.80. This correlation is depicted in Figure A.1. in the online appendix.

[9] The SSR data is available for the US, the UK, Japan, New Zealand, Australia, Canada, Switzerland, and the Eurozone economies.

[10] For all HP-filtered trends, we use the standard smoothing parameter $\lambda = 100$ for annual data, following Hodrick and Prescott (1997).

[11] In the online Appendix, we alternatively use the level of the inflation rate instead of the inflation gap as a robustness test; the results remain largely unchanged.

activity, as well as the real effective exchange rate (REER), which reflects external adjustment and exchange-rate transmission channels. Both CPI inflation and GDP are sourced from the World Bank, as is the REER. To capture the broader macroeconomic and financial transmission of monetary policy, the vector of endogenous variables is augmented with additional indicators, including government final consumption expenditure as a percentage of GDP, the unemployment rate, an interest rate spread defined as the difference between lending and deposit rates,[12] the total value of stocks traded as a share of GDP as our stock price index, and house prices.[13] Following the monetary economics literature, all variables enter the model in logarithms, except those already expressed as percentages, which are included in levels (Christiano et al., 1999; Jarociński and Karadi, 2020; Miranda-Agrippino and Rey, 2020).

2.2. Empirical model

We estimate an interacted panel VAR (IP-VAR) on a panel of 76 countries to examine how IT shapes the transmission of expansionary monetary policy shocks to income inequality. The empirical framework follows Sá et al. (2014) and allows the dynamic relationships among macroeconomic and financial variables to vary systematically with the monetary policy regime. The structural form of the model is given by:

$$J_{i,t}Y_{i,t} = \sum_{j=1}^{N} \tilde{C}_j D_{j,i} + \sum_{j=1}^{N}\sum_{k=1}^{L} \tilde{A}_{j,k} D_{j,i} Y_{i,t-k} + \tilde{C}^1 X_{i,t} + \sum_{k=1}^{L} \tilde{A}_k^1 X_{i,t} Y_{i,t-k} + \tilde{u}_{i,t}. \tag{1}$$

where $t = 1, \dots, T$ denotes time and $i = 1, \dots, N$ indexes countries. $Y_{i,t}$ is a $q \times 1$ vector of endogenous variables, as described above. The vector $\tilde{C}_j$ is a $q \times 1$ vector of country-specific intercepts. The vector of residuals, $\tilde{u}_{i,t}$, is assumed to be uncorrelated across countries and normally distributed with mean zero and a constant $q \times q$ covariance matrix $\tilde{\Sigma}$. The variable $D_{j,i}$ is a country-specific indicator that allows the VAR coefficients to vary across countries, capturing time-invariant country-level heterogeneity. This specification addresses the inconsistency of standard fixed-effects estimators in dynamic panels when slope coefficients are heterogeneous (Pesaran and Smith, 1995). Finally, $\tilde{A}_{j,k}$ denotes a $q \times q$ matrix of autoregressive coefficients for lag $k = 1, \dots, L$.

The interaction term $X_{i,t}$ influences the dynamic relationship between the endogenous variables through the coefficients $\tilde{A}_k^1$, and may also affect their levels through $\tilde{C}^1$. In our application, $X_{i,t}$ captures the monetary policy framework, specifically the adoption of an inflation-targeting regime, which is treated as a structural characteristic that conditions the transmission of monetary policy shocks over time. Time variation in the model therefore arises from changes in this institutional feature, rather than from stochastic time-variation in the coefficients, as in single-country VARs (e.g. Cogley and Sargent (2005), Primiceri (2005) and Canova et al. (2007)).

The $q \times q$ matrix $J_{i,t}$ is lower triangular with ones on the main diagonal, implying a recursive contemporaneous structure under which the covariance matrix of the structural shocks, $\tilde{\Sigma}$, is diagonal. For $q < w$, the contemporaneous effect of the qth-ordered variable on the wth-ordered variable is given by the (w, q) element of $-J_{i,t}$, which is modeled as:

$$J_{i,t}(w,q) = \sum_{j=1}^{N} J_j(w,q) D_{j,i} + J^1(w,q) X_{i,t}.$$

The coefficients $J_j(w,q)$ and $J^1(w,q)$ capture, respectively, country-specific baseline effects and the marginal impact of the IT regime. By construction, $J_{i,t}(w,q) = 1$ for $q = w$ and $J_{i,t}(w,q) = 0$ for $q > w$, reflecting the recursive ordering of the system, whereby variables ordered earlier do not respond contemporaneously to variables ordered later.

The reduced-form representation is obtained by premultiplying the structural system by $J_{i,t}^{-1}$. Note that the vector of reduced-form residuals, $u_{i,t}$, satisfies the condition $u_{i,t} = J_{i,t}^{-1}\tilde{u}_{i,t}$ and the Cholesky decomposition of the reduced-form covariance matrix $\Sigma_{i,t}$ is given by $\Sigma_{i,t} = V_{i,t}V_{i,t}'$, with $V_{i,t} = J_{i,t}^{-1}\tilde{\Sigma}^{1/2}$.

The recursive ordering reflects standard timing assumptions in the monetary policy literature. Government spending, GDP, the unemployment rate, inflation, and income inequality are ordered before the policy interest rate, while the interest rate spread, the real effective exchange rate, the stock market index, and house prices are ordered after the policy rate. This structure implies that macroeconomic aggregates respond to monetary policy shocks with a lag, while financial variables can react contemporaneously. At the same time, the ordering is consistent with the view that central banks condition their policy decisions on contemporaneous information about output, unemployment, and inflation within the year. Note that our main result remains robust when monetary policy shocks are identified using sign restrictions following Uhlig (2005).

IRFs are computed separately for the IT and non-IT regimes. To characterize the responses of a representative country, we set the country indicator variables to $D_{j,i} = 1/N$ for all j and evaluate the model at $X_{i,t} = 1$ for the IT regime and $X_{i,t} = 0$ for the non-IT regime. The resulting regime-specific structural coefficients are converted to reduced-form coefficients by premultiplying by the inverse of the contemporaneous impact matrix. IRFs are then obtained from the corresponding reduced-form VAR and interpreted as the dynamic responses of a typical country under each monetary policy regime (IT versus non-IT).

The dimensionality of our baseline VAR is non-trivial due to the inclusion of multiple macroeconomic and financial variables, interaction terms, and the panel structure. In such settings, standard OLS estimation can lead to overfitting and unstable inference,

[12] The interest rate spread is defined as the interest rate charged by banks on loans to private-sector customers minus the interest rate paid on demand, time, or savings deposits.

[13] All variables are obtained from the World Bank, except for house prices, which are sourced from the OECD and are therefore available only for OECD countries.

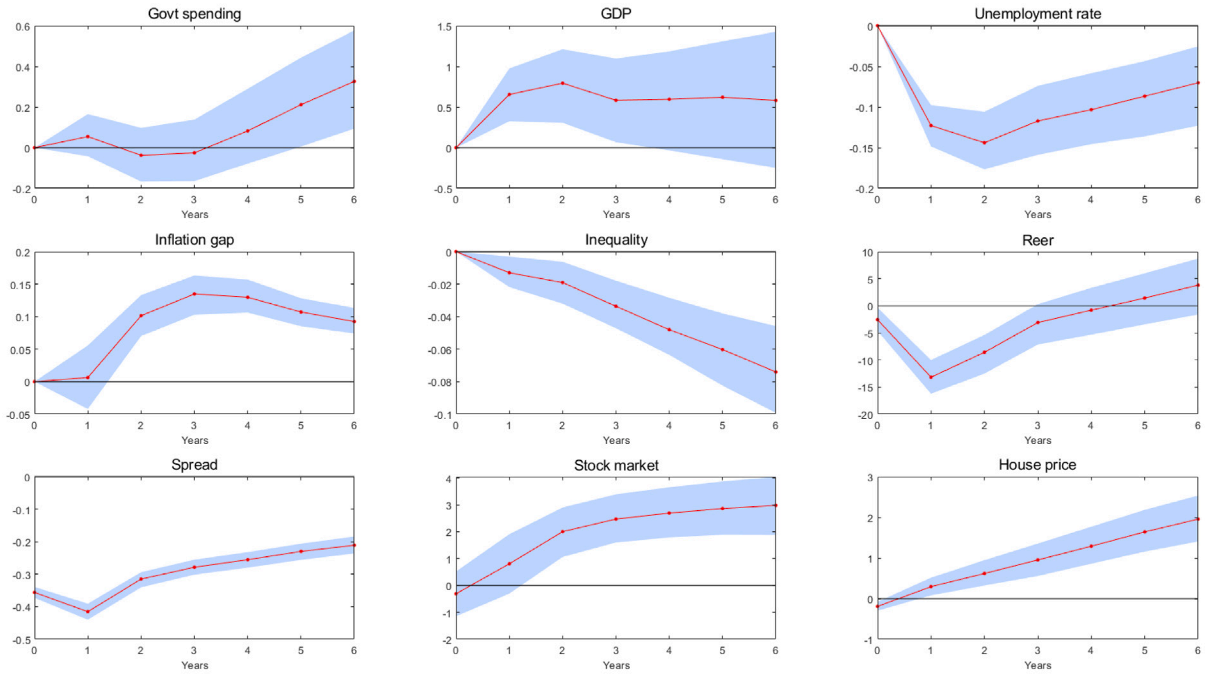


Fig. 2. IRFs from a Bayesian interacted panel VAR: IT regime.
Note: The figure shows impulse response functions from a Bayesian interacted panel VAR for IT regimes over the period 1980–2017. The shock corresponds to an exogenous 100 basis points reduction in the short-term interest rate. The red line denotes the posterior median response, and the blue shaded areas represent 68% intervals. Responses are reported over a six-year horizon. (For interpretation of the references to color in this figure legend, the reader is referred to the web version of this article.)

thereby weakening the reliability of structural analysis (Bańbura et al., 2010; Giannone et al., 2015). To address these concerns, we estimate the model using Bayesian methods with a Normal–Wishart prior, following Negro and Schorfheide (2011). This prior provides moderate shrinkage while allowing the data to play a central role in determining the posterior estimates and has been widely used in VAR applications. A similar Bayesian framework is adopted by Sá et al. (2014) in an interacted panel VAR setting. The model is estimated with two lags, consistent with Calza et al. (2013) and Sá et al. (2014).

3. Results

3.1. Baseline model

Our baseline VAR model is estimated using data for 27 IT regimes and 49 non-IT regimes over the period 1980–2017. Fig. 2 presents the results for the IT regime over a six-year horizon for all macroeconomic variables following a 100 basis points (i.e. one percentage point) reduction to the short-term interest rate.[14] The red line represents the median estimate, while the blue shaded areas depict the 68 credible intervals[15] The corresponding results for non-IT countries are reported in Fig. 3.

We begin by focusing on the response of income inequality measured by the net income Gini coefficient.[16] For countries operating under IT regimes, an expansionary monetary policy shock leads to a statistically significant and persistent reduction in income inequality. Quantitatively, a 100 basis-point expansionary monetary policy shock is associated with a 0.08 Gini-point decline in income inequality. By contrast, in non-IT regimes, the response of income inequality differs in direction. While income inequality increases initially following an expansionary monetary policy shock, this positive response is statistically significant only at the very short horizon and quickly becomes indistinguishable from zero thereafter. The absence of persistence and the limited horizon over which this effect appears sharply distinguish non-IT regimes from IT regimes, where inequality declines in a sustained manner. As shown in the following section, a plausible interpretation is that, in non-IT regimes, the financial channel of monetary policy may

[14] The response of the interest rate is shown in Figure A.1 of the online Appendix.

[15] Note that in contrast to the frequentist approach, it is standard practice in Bayesian VARs to report 68% error bands (Sims and Zha, 1999; Bańbura et al., 2010; Liu et al., 2014; Inoue and Rossi, 2019).

[16] Following Furceri et al. (2018), we focus on net income in the baseline analysis for two reasons. First, the SWIID dataset used in this study is originally constructed using net incomes, which reduces potential measurement error. Second, net income captures the overall distributional effects of monetary policy, including redistribution through taxes and transfers.

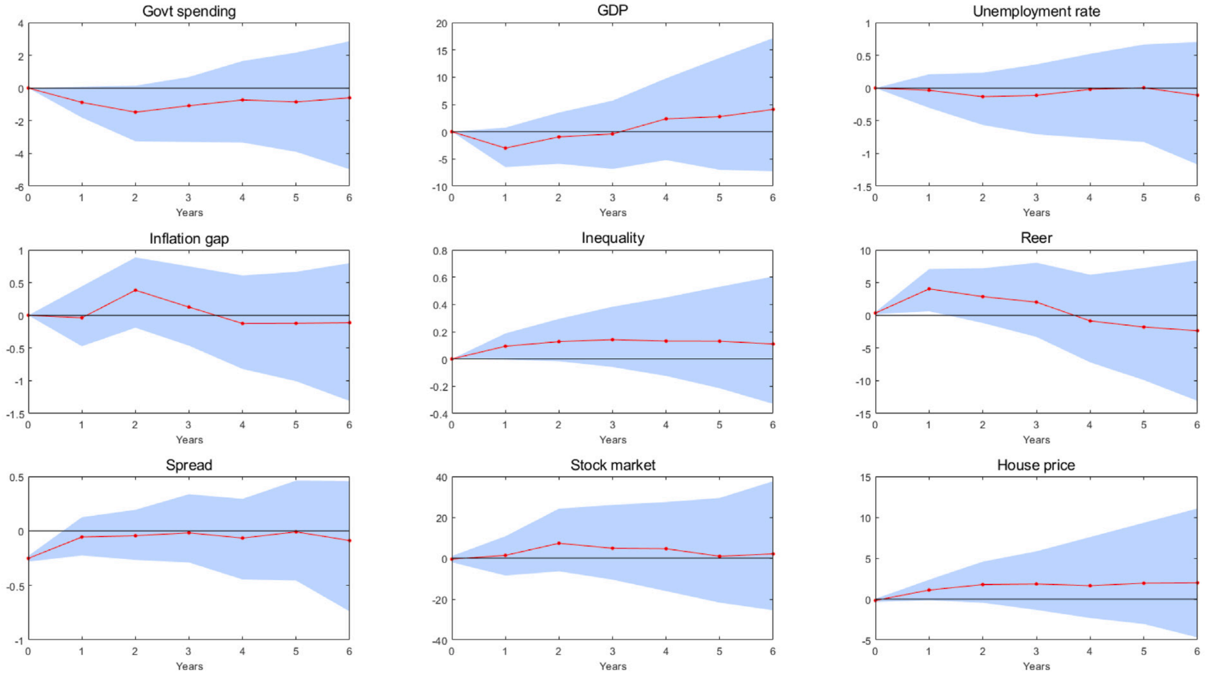


Fig. 3. IRFs from a Bayesian interacted panel VAR: non-IT regime.
Note: The figure shows impulse response functions from a Bayesian interacted panel VAR for non-IT regimes over the period 1980–2017. The shock corresponds to an exogenous 100 basis points reduction in the short-term interest rate. The red line denotes the posterior median response, and the blue shaded areas represent 68% credible intervals. Responses are reported over a six-year horizon. (For interpretation of the references to color in this figure legend, the reader is referred to the web version of this article.)

dominate in the short run, generating temporary fluctuations in asset prices and leading to a transient increase in inequality (O'Farrell and Rawdanowicz, 2017).

These results speak directly to the broader debate on the distributional consequences of monetary policy. First, the countercyclical response of income inequality echoes the results of Coibion et al. (2017), who study the United States and show that the strongest redistributive effects of monetary policy occurred during the Volcker disinflation period, which was characterized by monetary tightening.[17] Our results are also consistent with evidence from other empirical studies. For instance, Mumtaz and Theophilopoulou (2017) document similar effects for the United Kingdom, while Samarina and Nguyen (2024), Guerello (2018), Meriküll and Rottner (2024) and Lenza and Slacalek (2024) provide supporting evidence for panels of European and OECD countries, showing that expansionary conventional monetary policy is associated with declining income inequality. Comparable findings are reported by Furceri et al. (2018) for a sample of 20 advanced and 12 emerging economies, the majority of which operate under IT regimes.[18]

Second, this short-lived positive response of income inequality in non-IT countries, statistically significant only in the initial period and quickly becoming indistinguishable from zero thereafter, is consistent with an earlier strand of the literature focusing on large samples of developed and developing economies during periods when inflation-targeting regimes had not yet been widely adopted. Using standard panel data techniques, rather than VAR or local projection methods, these studies typically find a weak or unstable relationship between expansionary monetary policy and income inequality, and in some cases document a positive short-run association (Bulir and Gulde, 1995; Romer and Romer, 1998; Easterly and Fischer, 2001).

Our findings contrast with studies focusing on Japan, where there is now broad empirical agreement that monetary easing has exacerbated income inequality (Saiki and Frost, 2014; Inui et al., 2017; Taghizadeh-Hesary et al., 2020). Japan, however, may represent a special case, reflecting substantial differences in household balance sheets and labor market institutions relative to other advanced economies (Saiki and Frost, 2018).[19]

Turning to the impulse responses of the macroeconomic variables following a decrease in the short-term interest rate, IT regimes exhibit dynamics that are consistent with standard theoretical predictions. Real GDP increases gradually and significantly, reaching

[17] During the Volcker period, the Federal Reserve operated in a manner akin to an inflation-targeting central bank (Goodfriend, 2004; Ball and Sheridan, 2004).

[18] When the authors distinguish between advanced and emerging economies, they find statistically similar effects.

[19] With respect to inflation targeting, Japan may be regarded as both a targeter and a non-targeter during our sample period, given its prolonged episode of deflation without the formal adoption of an inflation-targeting framework. Had a formal IT regime been implemented, inflation declines might have been more limited, potentially strengthening the macroeconomic and distributional effects of monetary policy (Ball and Sheridan, 2004).

a peak response of around 0.75 percent after approximately two years. The inflation gap responds positively, peaking within three years and subsequently reverting, consistent with a temporary overshooting of the inflation target.[20] In contrast, non-IT countries display largely muted and statistically insignificant responses of both GDP and the inflation gap, as illustrated in Fig. 3. These findings are consistent with the view that inflation-targeting regimes are more successful in stabilizing inflation dynamics while supporting real economic activity, as emphasized in Svensson (2010) and supported by empirical evidence in Cecchetti and Ehrmann (1999), Neumann and Von Hagen (2002), Mishkin and Schmidt-Hebbel (2007), Creel and Hubert (2015) and Gallegos (2023), though they contrast with studies that find limited real effects of IT adoption (Lin and Ye, 2007).

The response of the REER also differs across monetary policy regimes. Under IT regimes, the REER exhibits a statistically significant decline on impact, consistent with a depreciation of the domestic currency following an expansionary monetary policy shock, in line with standard theoretical mechanisms and empirical findings for advanced economies (Kim and Roubini, 2000; Faust and Rogers, 2003; Scholl and Uhlig, 2008; Inoue and Rossi, 2019). By contrast, under non-IT regimes, the REER response is positive, suggesting an appreciation. This pattern likely reflects the prevalence of managed or fixed exchange rate arrangements among non-IT countries, which can actively influence exchange rate dynamics through foreign exchange market interventions.[21]

Finally, the responses of additional macro-financial variables in IT countries broadly align with those typically observed following an expansionary monetary policy shock. The unemployment rate declines, credit spreads narrow on impact, equity prices and house prices increase gradually over time. Government spending remains subdued reflecting the absence of a direct monetary policy effect, but turns positive at longer horizons. This delayed response is consistent with lower interest rates reducing the cost of public borrowing, thereby expanding fiscal space and enabling higher government expenditure in the medium to long run.

Overall, our comparative evidence on the impact of the monetary policy shock on income inequality for IT countries relative to non-IT countries suggests that implementing IT leads to significant reductions in the Gini coefficient for income inequality. In Section 3.3, we examine the specific transmission channels underlying these effects.

3.2. Robustness checks

We conduct an extensive set of robustness checks to assess the sensitivity of our baseline results to alternative specifications, identification strategies, and sample definitions. Across all exercises, our central finding remains unchanged: expansionary monetary policy shocks reduce income inequality under IT regimes, while the corresponding effects under non-IT regimes are weak, short-lived, or statistically insignificant. A brief overview of the robustness exercises is provided below, while each robustness check is reported in detail in Section 1.3 of the Online Appendix.

First, we examine whether the results are driven by major macroeconomic episodes or changes in the conduct of monetary policy. Allowing for different responses before and after the Global Financial Crisis (GFC), and accounting for the zero lower bound through the use of shadow short rates in economies that adopted unconventional policies, leaves the baseline inequality results under IT intact. Second, we verify that our findings are not sensitive to the classification of inflation-targeting regimes. Re-estimating the model using progressively broader definitions of IT, incorporating recent adopters, quasi-targeters, and euro-area members, yields quantitatively similar inequality responses across all specifications. Third, we assess distributional heterogeneity by replacing the Gini coefficient with income percentile ratios. Specifically, we consider the 90/10 and 99/10 income ratios, which capture income dispersion across different segments of the distribution. Under IT, both ratios decline following an expansionary monetary policy shock, consistent with the reduction in overall income inequality and indicating stronger effects at the middle and lower segments of the distribution. Finally, we consider an alternative identification strategy based on sign restrictions. The resulting inequality responses closely mirror those obtained under recursive identification, confirming that the inequality-reducing effects of monetary policy under IT are robust to identification choices.

3.3. Transmission channels

Our baseline results indicate that expansionary monetary policy tends to reduce income inequality in countries operating under IT monetary policy frameworks. This section investigates the mechanisms through which this effect materializes, with a particular focus on the employment and financial channels.

The employment channel plays a key role in reducing income inequality, as expansionary monetary policy stimulates job creation, particularly benefiting low-skilled and low-income households who rely heavily on wage income. This effect is further reinforced under effective IT regimes, whose institutional features, such as improved central bank credibility, anchored inflation expectations, increased transparency, and strengthened accountability (Bernanke, 2003; Ascari et al., 2017; Cole and Martínez-García, 2023), foster a macroeconomic environment characterized by lower average inflation and reduced inflation volatility (Lin and Ye, 2009; Batini and Laxton, 2007; Hyvonen et al., 2004; Vega and Winkelried, 2005; Mishkin and Schmidt-Hebbel, 2007; Pétursson et al., 2009), as well as enhanced resilience to large shocks (Duncan et al., 2022). These conditions, in turn, promote investment and employment growth that more strongly benefit lower-income households.

[20] Temporary price-level overshooting has been documented in the literature and may reflect either deliberate policy accommodation in response to macroeconomic contingencies or imperfect inflation-targeting performance (Romer and Romer, 2007).

[21] More than 95% of economies classified as currency unions, pre-announced pegs, currency boards, de facto pegs, or horizontal band regimes according to Ilzetzki et al. (2022) fall into the non-IT category.

In contrast, the financial channel tends to increase income inequality following expansionary monetary policy. Rising asset prices generate capital gains that accrue primarily to wealthier households holding financial assets, while inflation erodes the nominal incomes of poorer households, especially those without inflation-indexed earnings (O'Farrell and Rawdanowicz, 2017; Samarina and Nguyen, 2024; Doepke and Schneider, 2006). However, the institutional features of IT regimes highlighted above, help cushion these adverse effects on lower-income groups (Bulir and Gulde, 1995; Bulíř, 2001; Easterly and Fischer, 2001; Romer and Romer, 1998; Galli and van Der Hoeven, 2001), thereby mitigating the inequality-increasing impact of the financial channel.[22]

To assess whether monetary policy impacts inequality through the employment and financial channels, we conduct two counterfactual experiments, one for each channel. Our counterfactuals are constructed as follows: we apply a monetary policy shock and then construct a linear combination of the remaining structural shocks such that the impulse response of the targeted transmission variable, either the unemployment rate (as a proxy for the employment channel) or the stock market index (as a proxy for the financial channel), is set to zero at each forecast horizon. This approach follows Mountford and Uhlig (2009) and Mumtaz and Theophilopoulou (2020), and addresses potential concerns related to the Lucas Critique, as it preserves the model's structure and parameters by adjusting only the composition of the remaining shocks required to offset the effect of the monetary policy shock on the targeted variable.

Given that we have established the significant impact that expansionary monetary policy has on reducing income inequality in IT countries, we therefore focus primarily on economies operating under an IT regime. The results are depicted in Fig. 4.

The counterfactual response of income inequality is depicted by the solid red line, together with its confidence interval, and is compared against the baseline response shown by the dotted line. When the employment channel is shut down, the counterfactual response becomes markedly more muted, and the baseline response clearly falls outside the confidence bands. This pronounced divergence between the counterfactual and baseline responses indicates that, for countries operating under an IT regime, the employment channel constitutes a key mechanism driving the reduction in income inequality. By contrast, the financial channel appears to play a more limited role. As shown in Fig. 4, the baseline response remains within the confidence bands of the counterfactual response obtained when the financial channel is shut down throughout the entire forecast horizon. Note that the baseline response lies above the counterfactual, suggesting that the financial channel dampens the decline in income inequality and therefore partially offsets the inequality-reducing effects operating through the employment channel. However, consistent with the discussion above, this dampening effect is small in magnitude and statistically insignificant. This finding is consistent with previous studies, such as Samarina and Nguyen (2024) and O'Farrell and Rawdanowicz (2017), which highlight the relatively limited role of the financial channel in the transmission of monetary policy.[23]

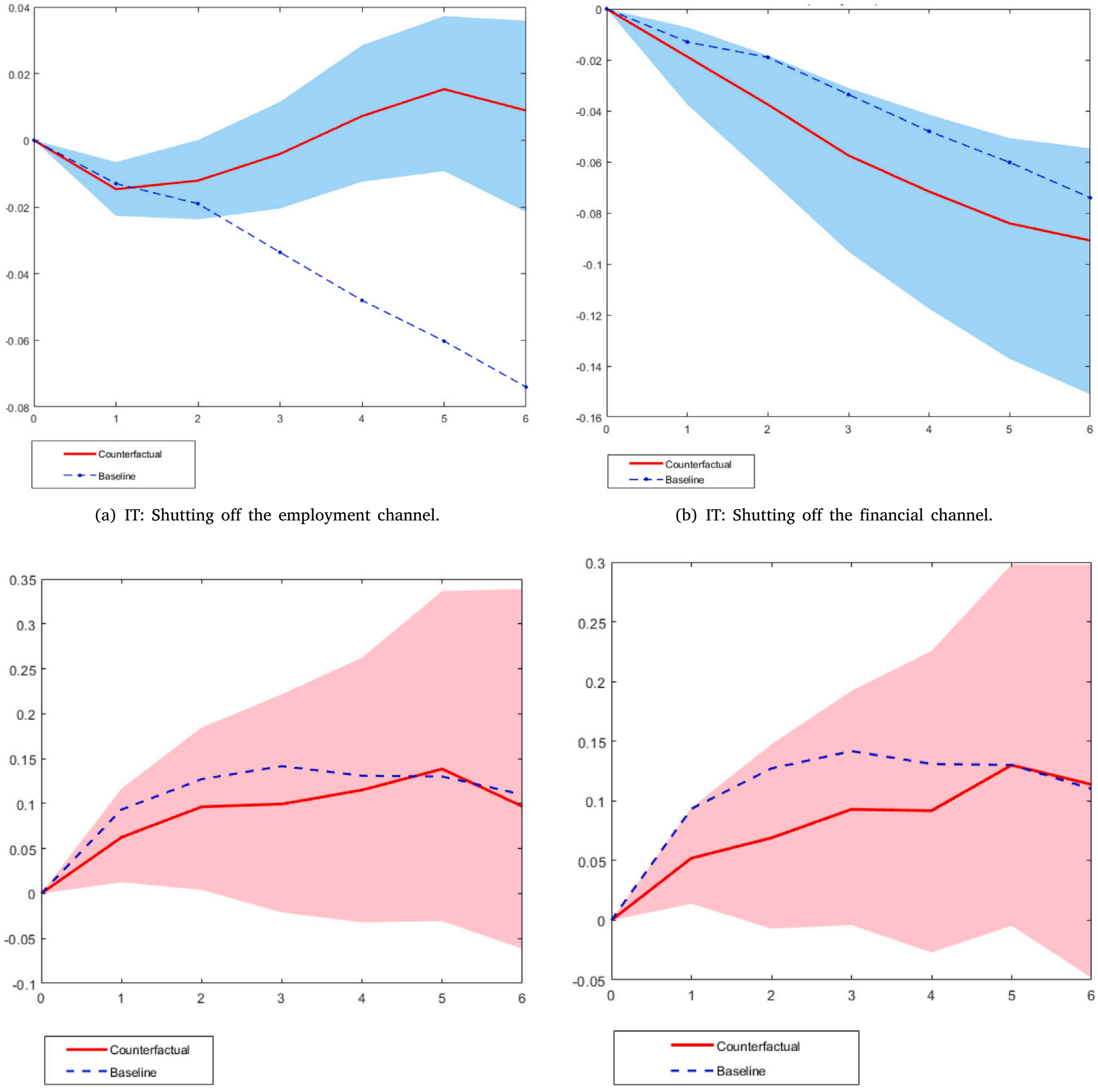
The counterfactual responses for non-IT regimes are reported in the lower panels of Fig. 4. The employment channel does not appear to play an important role in the transmission of monetary policy, as the baseline response of income inequality remains largely within the confidence bands of the counterfactual response. Note, however, that the fact that the baseline response lies above the counterfactual indicates that the employment channel fails to generate inequality-reducing effects in non-IT regimes and is instead associated with a weak increase in income inequality. One possible explanation is that, in the absence of a credible IT framework, employment gains following an expansionary monetary policy shock are limited or unevenly distributed, thereby failing to deliver broad-based income improvements for lower-income households. Turning to the financial channel, note that at very short horizons, the baseline response lies marginally outside the upper band of the counterfactual response. This pattern points to a non-negligible role for the financial channel in driving the short-run increase in income inequality following an expansionary monetary policy shock in non-IT regimes. Although these effects are short-lived and lose statistical significance at longer horizons, they suggest that financial market dynamics are the primary driver of the temporary rise in income inequality observed in the baseline response in Fig. 3.

3.4. Institutional frameworks and inflation environments

IT frameworks are often accompanied by institutional features that can influence the implementation and transmission of monetary policy (Crowe, 2010; Alpanda and Honig, 2014). This section examines whether the distributional effects of monetary policy that we attribute to IT instead reflect broader institutional characteristics or macroeconomic environments that frequently accompany IT adoption. We focus on central bank independence (CBI), central bank transparency (CBT), and prevailing inflation regimes, distinguishing between high- and low-inflation periods, to assess whether IT has an effect on inequality over and above these related factors.

[22] For clarity, the greater role of the employment channel in IT regimes is not due to IT disproportionately strengthening it relative to the financial channel. Rather, as explained in the main text, IT improves the macroeconomic environment in a manner that benefits both channels. However, since the employment channel reduces inequality and the financial channel increases it in response to an expansionary monetary policy shock, the employment channel's inequality-reducing impact is amplified, while the financial channel's inequality-increasing effect is dampened.

[23] As an alternative to the stock market index, we use house prices to capture the housing price channel. The results mirror those for the financial channel: the baseline response of inequality remains well within the counterfactual error bands, indicating that housing prices play a limited role in driving income inequality (see Online Appendix Figure A.8.). One notable difference is that the housing channel operates in the same direction as the employment channel, contributing to a decline in inequality. This is consistent with the broader distribution of real estate wealth relative to financial assets, which are concentrated among households at the top of the wealth distribution (Haliassos and Bertaut, 1995; Evgenidis and Fasianos, 2021; Fasianos and Tsoukalis, 2023). As a result, income returns from housing generate a more equalizing response to monetary policy shocks, whereas returns from financial assets accrue disproportionately to wealthier households.



(a) IT: Shutting off the employment channel.

(b) IT: Shutting off the financial channel.

(c) Non-IT: Shutting off the employment channel.

(d) Non-IT: Shutting off the financial channel.

Fig. 4. Counterfactual transmission channels under IT and non-IT regimes.
Note: The top panels show counterfactual results of income inequality responses under the IT regime, while the bottom panels report the same channel shutdown counterfactuals for non-IT regimes. Each panel isolates a specific transmission channel as described in the empirical section. (For interpretation of the references to color in this figure legend, the reader is referred to the web version of this article.)

3.4.1. Central bank independence

Our results so far indicate that IT improves macroeconomic performance and mitigates income inequality following expansionary monetary policy shocks. This naturally raises the question of whether CBI plays an additional role in shaping the transmission of monetary policy to inequality and the broader macroeconomy, particularly within IT regimes.

CBI is often viewed as a key institutional feature that can enhance the credibility and effectiveness of monetary policy by anchoring inflation expectations and insulating policy decisions from short-term political pressures (Mishkin, 2000; Alpanda and Honig, 2014). In principle, higher levels of CBI may therefore strengthen the transmission of monetary policy under IT by ensuring a sustained commitment to price stability. However, the existing literature does not provide a definitive answer as to whether CBI reinforces the effects of inflation targeting or merely reflects a complementary institutional characteristic. For instance, Svensson

(1997) and Bernanke et al. (2000) argue that countries with lower initial levels of central bank independence may experience larger gains from adopting IT, suggesting that the interaction between IT and CBI is not straightforward.[24]

To assess the joint role of these institutional features, we extend our baseline specification by incorporating both the IT indicator and the CBI dummy as interaction variables. This allows us to examine whether the distributional effect of monetary policy shocks under IT depends on the degree of independence, and whether CBI on its own generates effects comparable to those observed under IT regimes.

We measure CBI using the index of Garriga (2016).[25] The original series provides scores combined in a single index, ranging from 0 (no independence) to 1 (highest independence), and runs from 1970 to 2012.[26] We embed the index into our IP-VAR framework as an exogenous dummy by constructing two different regimes of independence: one for countries scoring "low CB independence" and one for countries scoring "high CB independence", which varies by country and across time. The observations are assigned to each threshold depending on whether their score is above or below the sample median.[27]

Fig. 5 examines whether CBI affects the distributional impact of monetary policy shocks. When CBI is considered in isolation, the response of income inequality to an expansionary monetary policy shock is largely statistically insignificant, regardless of whether central bank independence is high or low. This suggests that institutional independence alone is not sufficient to generate meaningful inequality-reducing effects. A different picture emerges once inflation targeting IT is taken into account. Conditioning jointly on the monetary policy regime and the degree of central bank independence reveals that the reduction in income inequality following an expansionary monetary policy shock is driven primarily by the presence of IT. In countries operating under IT, income inequality declines significantly after the shock, both in low- and high-CBI environments. By contrast, in non-IT regimes, the response of inequality remains weak and statistically insignificant, and in some cases points to a short-lived increase in inequality at very short horizons. Importantly, while IT is the key condition for generating inequality-reducing effects, a higher degree of central bank independence modestly reinforces these gains: within IT regimes, the decline in income inequality is slightly more pronounced when central bank independence is high compared to low.

3.4.2. *Central bank transparency*

Central bank transparency is often viewed as an important component of inflation targeting, as IT regimes typically place greater emphasis on communication, forward guidance, and the publication of forecasts (Geraats, 2002; Chortareas et al., 2002; Walsh, 2009). This subsection examines whether transparency itself conditions the distributional effects of monetary policy shocks, or whether its role is primarily complementary to IT.

We measure CBT using the index developed by Dincer et al. (2022). The index is widely recognized in the literature and covers more than 100 economies. It ranges from 0 (no transparency) to 15 (high transparency) and can be disaggregated into five sub-components, each reflecting a different dimension of CBT.[28] We construct two different regimes of CBT: "low CBT" and "high CBT". Observations are assigned to each regime depending on whether their score is above or below the median of the full sample, which is estimated at 4.5.[29]

Fig. 6 shows that, when transparency is considered in isolation, the response of income inequality to an expansionary monetary policy shock remains largely statistically insignificant in both high- and low-CBT regimes. This indicates that CBT alone does not generate inequality-reducing effects, mirroring the evidence obtained for central bank independence. Conditioning jointly on inflation targeting and transparency yields a clearer pattern. In countries operating under IT, expansionary monetary policy shocks are associated with a significant decline in income inequality, regardless of the level of transparency. The reduction in the Gini coefficient is more pronounced and slightly more persistent in high-CBT IT regimes than in low-CBT IT regimes. Although these differences are quantitatively modest, they suggest that CBT plays a supportive role by reinforcing the effectiveness of inflation targeting, rather than acting as an independent driver of distributional outcomes.

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24 See also Rogoff (2021) for a discussion of the role of institutional credibility in anchoring inflation expectations.

25 The author expands the original classification of Cukierman (1992) by including 159 non-OECD economies and accounts for a variety of CBI characteristics, including the status of the chief executive officer of the bank (appointment, dismissal, and term of office), the bank's policy formulation, CB objectives, and restraints from lending to the government.

26 To cover the period beyond 2012, we extrapolate the dataset up to 2017 using the latest scores available for each country, assuming no variation in CB independence after 2012. This is not too strong an assumption considering that our thresholds include only two cases, that of high and low CBI, and that the degree of CBI is slow-moving over time.

27 López-Villavicencio and Pourroy (2022) use a similar transformation of the CBI index.

28 The first component measures the degree of openness regarding monetary policy objectives; the second component focuses on the economic information used by monetary policy makers, including data and CB publications; the third component captures the presence of an explicit monetary policy rule or strategy, as well as disclosures related to the decision-making process; the fourth component assesses policy transparency, including the prompt disclosure of policy decisions, complementary explanations, and potential indications of future policy actions; the fifth component, operational transparency, examines the disclosure of evaluations of deviations from the main monetary policy targets and outcomes, as well as discussions of deviations affecting monetary policy transmission.

29 As an imputation strategy for missing data before 1998 in some countries, we assign the maximum transparency score to those countries with missing observations that operate under a fixed exchange rate regime. For the rest of the countries, we extrapolate backwards using the oldest scores available up to 1980 for each country.

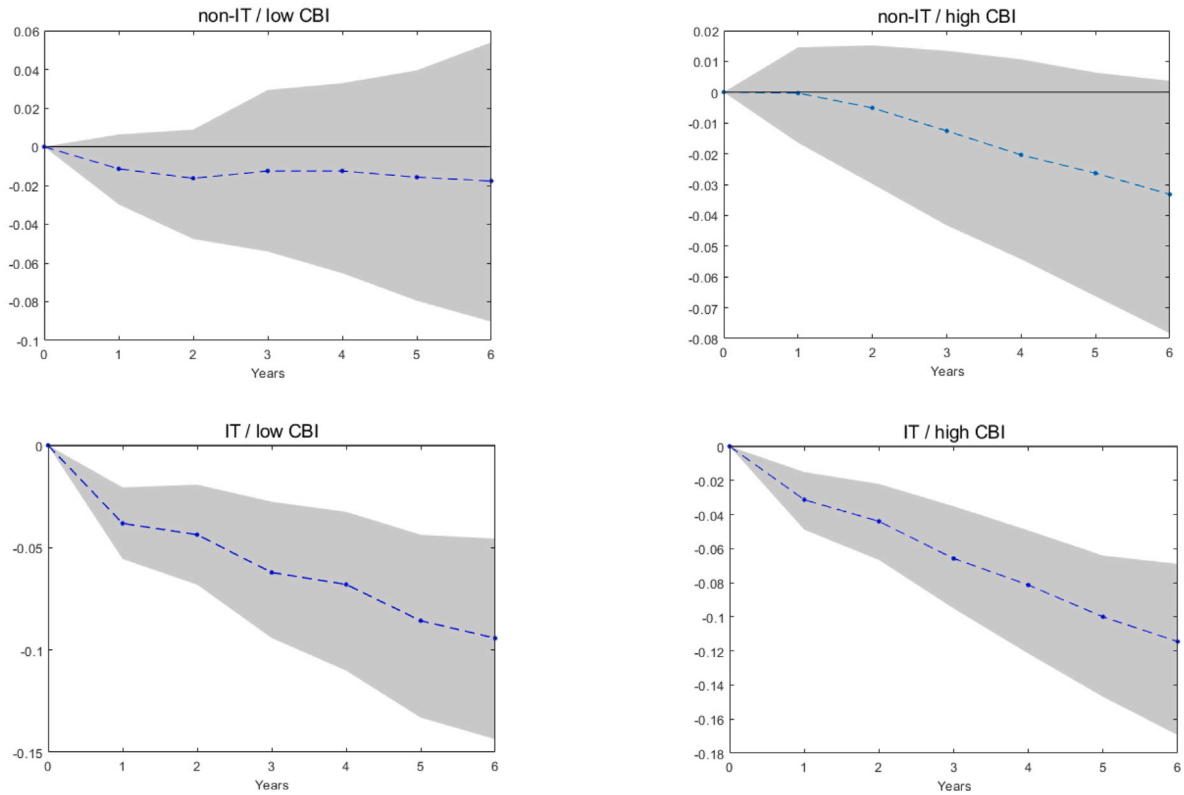

Fig. 5. Responses of income inequality under IT vs. non-IT regimes and high vs. low CBI.
Note: Impulse responses of income inequality to an expansionary monetary policy shock across monetary policy regimes and levels of central bank independence (CBI). Solid lines denote point estimates, while shaded areas represent the credible intervals.

3.4.3. Inflation environments

A potential concern is whether our results reflect periods of low and stable inflation, rather than the monetary policy framework itself. Many advanced economies, such as the US, experienced prolonged episodes of low inflation during the sample period, even in the absence of a formal IT regime. Conversely, some countries, such as Turkey, maintained an explicit IT framework while facing persistently high inflation. To isolate the role of IT from the prevailing inflation environment, we examine whether our baseline findings hold when distinguishing between high- and low-inflation regimes across both IT and non-IT countries.

We define inflation regimes using a common threshold applied uniformly across countries. Specifically, a period is classified as high inflation when CPI inflation exceeds the median inflation rate of the full sample, and as low inflation otherwise. The median inflation rate in our sample is estimated at 4.09 percent, which is very close to 4 percent, a cutoff commonly used in the literature to mark a change in agents' attention and expectation formation in higher-inflation environments (e.g., Weber et al. (2024)).[30] This inflation-regime dummy is then interacted with the IT indicator.

As in the baseline results, inequality responses are measured as changes in the Gini coefficient in original units. The results are depicted in Fig. 7. Under IT regimes, the monetary policy shock leads to a persistent decline in income inequality in both low- and high-inflation periods. The magnitude and persistence of the inequality response are very similar across the two inflation regimes, with only modest differences at longer horizons and largely overlapping confidence intervals throughout the forecast horizon. This indicates that, conditional on the presence of IT, changes in the inflation regime do not materially alter the transmission of monetary policy to inequality.

In contrast, when IT is not in place, the response of income inequality is weak and statistically insignificant in both low- and high-inflation environments. While point estimates differ somewhat across inflation regimes in non-IT countries, being slightly more negative in low-inflation periods and closer to zero or mildly positive in high-inflation periods, these effects are not statistically significant over the majority of the forecast horizon.

[30] To further assess robustness, we perform the same exercise using alternative absolute cutoffs (4.5 and 5 percent). We also consider an alternative specification in which inflation regimes are defined using country-specific thresholds based on deviations from each country's mean inflation. These additional results confirm the stability of our conclusions. These robustness checks confirm our results and are available upon request.

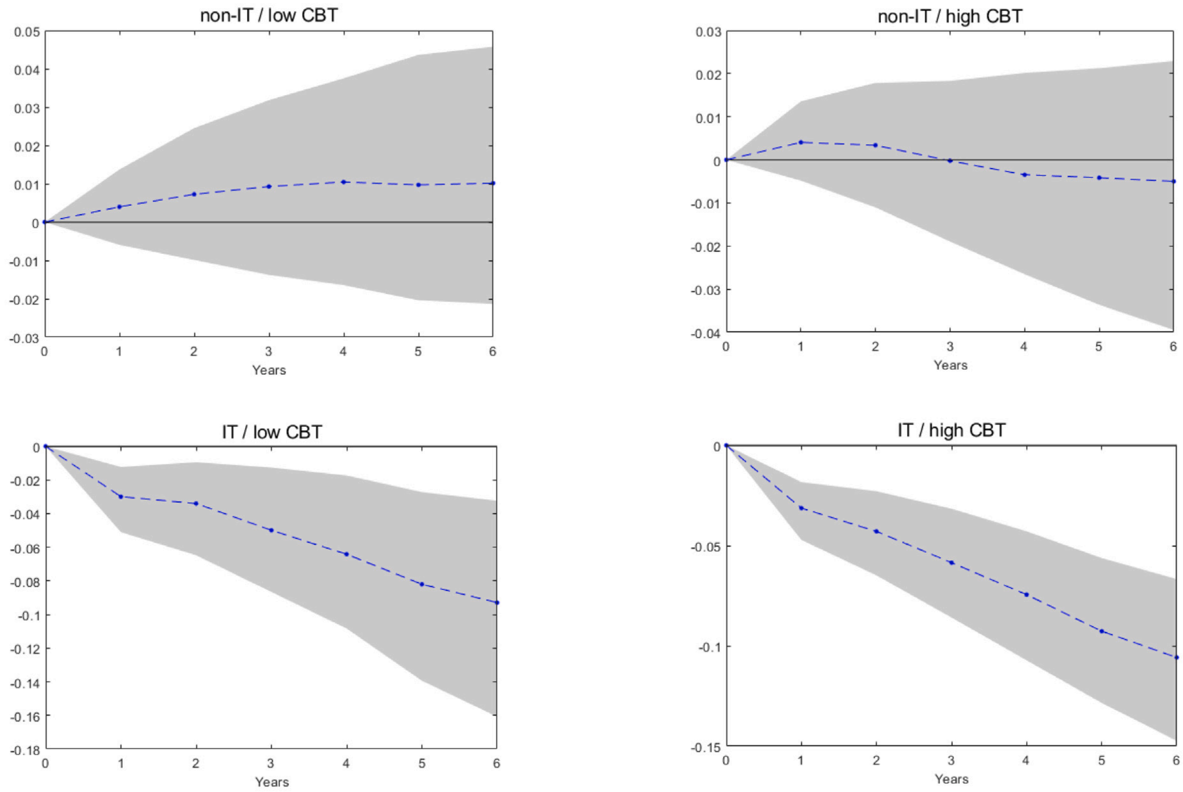

Fig. 6. Responses of income inequality under IT vs. non-IT regimes and high vs. low CBT.
Note: Impulse responses of income inequality to an expansionary monetary policy shock across monetary policy regimes and levels of central bank transparency (CBT). Solid lines denote point estimates, while shaded areas represent confidence intervals.

4. Monetary policy pass-through to inequality: PSM evidence

To address potential selection bias in the adoption of IT, we also employ a PSM approach (Rosenbaum and Rubin, 1983; Rubin and Thomas, 1996). To measure differences in inequality between IT and non-IT countries using the PSM, we derive an outcome variable reflecting the pass-through from monetary policy to inequality. We estimate the following pass-through specification, which is based on the inequality equation in the VAR specification but modified in line with the relevant pass-through literature (Campa and Goldberg, 2005; López-Villavicencio and Pourroy, 2019, 2022):

$$\Delta g_{i,t} = \alpha + \sum_{k=1}^{K} \theta_{i,t}^{(k)} \Delta r_{i,t-k} + \rho \Delta Y_{i,t} + \lambda \Delta \pi_{i,t} + \delta \Delta e_{i,t} + \epsilon_{i,t} \tag{2}$$

where $\Delta g_{i,t}$ denotes the difference in the Gini coefficient from period $t-1$ to period t in country i.[31] $\Delta Y_{i,t}$ represents the difference in GDP, $\Delta \pi_{i,t}$ denotes the difference in the CPI, and $\Delta e_{i,t}$ represents the change in the REER. The coefficient of interest, $\sum_{k=1}^{K} \theta_{i,t}^{(k)}$, represents the total pass-through from monetary policy to inequality, capturing the cumulative impact of lagged changes in the interest rate ($\Delta r_{i,t-k}$) on changes in inequality.[32] Each time-varying lag coefficient $\theta_{i,t}^{(k)}$ is estimated by specifying the following shift equation for each lag k: $\theta_{i,t}^{(k)} = \theta_{i,t-1}^{(k)} + c v_{i,t}^{(k)}$, where each pass-through parameter $\theta^{(k)}$ depends on an autoregressive term and $v_{i,t}^{(k)} \sim \mathcal{N}(0, Q_{i,t})$.[33] Consistent with the VAR model, we include two lags of the interest rate ($K = 2$) to appropriately capture delayed effects and inertia in the monetary policy transmission process. Using the resulting time-varying total pass-through coefficient $\sum_{k=1}^{2} \theta_{i,t}^{(k)}$, we assess the treatment effect of IT adoption on this coefficient. The estimation of interest can be expressed using the

[31] Since the Gini coefficient is measured on a 0–1 scale, the estimated pass-through effects should be interpreted as level changes (i.e., Gini-point equivalents), not percentage changes.

[32] Note that the pass-through equation is specified in first differences as the Gini coefficients in the SWIID dataset and the other underlying series are found to be integrated of order one (Christopoulos and McAdam, 2017; López-Villavicencio and Pourroy, 2019).

[33] Note that the above equations represent a state space model. These models are estimated using the Kalman filter to recursively obtain linear, least-squares forecasts of unknown coefficients conditional on past information (López-Villavicencio and Pourroy, 2019, 2022).

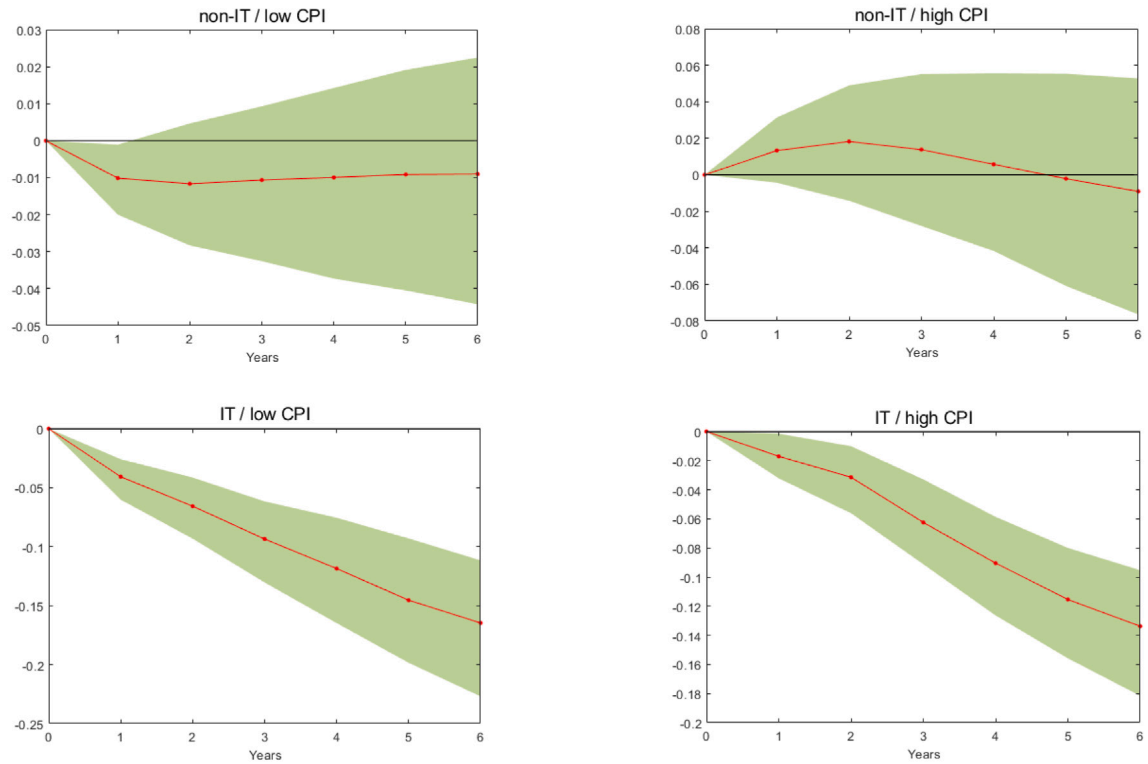

Fig. 7. Inequality responses: IT vs. non-IT and high vs. low inflation regimes.
Note: The figure reports impulse responses of income inequality to an expansionary monetary policy shock, distinguishing between IT and non-IT regimes, as well as high- and low-inflation periods. High- and low-inflation regimes are defined using a common threshold across countries based on the sample median inflation rate.

following average treatment on the treated (ATT) equation:

$$ATT = E[Y_{i,1}|D_i = 1] - E[Y_{i,0}|D_i = 1] \tag{3}$$

where D is the IT adoption dummy, with $\theta_{i,t}$ now being set as the outcome variable Y. $E[Y_{i,0}|D_i = 1]$ represents the counterfactual outcome that would have been observed if the IT country had not adopted IT, while $E[Y_{i,1}|D_i = 1]$ is the outcome observed once IT is adopted. The challenge in this estimation is that the counterfactual $E[Y_{i,0}|D_i = 1]$ is not observable. To address it, we rely on PSM methods and follow the steps outlined below.

First, we estimate the propensity score, i.e., the probability that a country adopts an inflation targeting (IT) regime conditional on observable characteristics.[34] We obtain the propensity score from a probit model in which the dependent variable equals one in the year a country switches to IT and zero otherwise. Consistent with the literature on the determinants of IT adoption, the model includes lagged inflation, broad money growth, lagged GDP, trade openness, the government current account balance, an index of exchange rate regime flexibility (Ilzetzki et al., 2022), and a measure of central bank independence (Garriga, 2016).[35]

Second, we proceed with matching treated observations (IT adoption) with control units (non-IT countries) based on the estimated propensity scores. We employ four commonly used matching algorithms, including nearest neighbor methods, kernel matching, local linear matching, and radius matching. These algorithms aim to find suitable control units for each treated observation based on their similarity in terms of the propensity scores.[36]

Third, after matching the observations, we examine the treatment effect by estimating the difference in the pass-through variable $\theta_{i,t}$ between the treated and control units. An additional assumption of the PSM requires comparable control observations for each treated observation ($p(X_i) < 1$). In other words, we need to ensure that the range of propensity scores for the treated observations

[34] This approach summarizes information from multiple covariates into a single index (Rosenbaum and Rubin, 1983).

[35] These variables are standard in empirical studies of IT adoption (Lin and Ye, 2007; Truman, 2003; López-Villavicencio and Pourroy, 2022).

[36] The underlying assumption of this methodology is the presence of conditional independence. This assumption implies that the potential outcomes (the outcome variable under treatment and under no treatment) are independent of the treatment assignment, conditional on a set of covariates X. In other words, once we control for the observed characteristics captured by X, the treatment assignment is independent of the potential outcomes.

Table 1
Average treatment effect of IT on the pass through of monetary policy to income Inequality.

	Nearest Neighbor (5) (1)	Nearest Neighbor (1) (2)	Kernel (3)	Local Linear (4)	Radius (0.05) (5)
Baseline PSM: Treatment is IT adoption					
ATT	0.015**	0.023*	0.017*	0.014*	0.017**
	(0.006)	(0.009)	(0.009)	(0.007)	(0.005)
# Treated/# Observations	446/1348	446/1348	446/1348	446/1348	446/1348
Alternative PSM specifications					
Exchange Rate Regime	0.018*	0.020*	0.015*	0.013	0.018**
	(0.010)	(0.010)	(0.007)	(0.011)	(0.008)
# Treated/# Observations	395/1199	395/1199	395/1199	395/1199	395/1199
Government CAB	0.013**	0.027*	0.014**	0.013*	0.015**
	(0.006)	(0.015)	(0.006)	(0.007)	(0.006)
# Treated/# Observations	374/1418	374/1418	374/1418	374/1418	374/1418
CB Independence	0.016**	0.014	0.010*	0.011	0.019*
	(0.007)	(0.009)	(0.006)	(0.008)	(0.010)
# Treated/# Observations	447/1650	447/1650	447/1650	447/1650	447/1650

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.
Note: Observed coefficient is the treatment effect (the difference between the treated and controls). When the pass through from monetary policy to inequality is higher for the treatment group (inflation targeters) than the controls, the observed coefficient is positive. Bootstrapped standard errors using 500 replications are presented in parentheses. Following the common support assumption we discard units whose estimated propensity scores are lower than the lowest score among the treated units.

overlaps with the range of propensity scores for the control observations. Therefore, we can express Eq. (3) in the following form:

$$ATT = E[Y_{i,1}|D_i = 1, p(X_i)] - E[Y_{i,0}|D_i = 0, p(X_i)] \tag{4}$$

where $E[Y_{i,1}|D_i = 1, p(X_i)]$ represents the outcome $\theta_{i,t}$ in countries that have implemented IT, and $E[Y_{i,0}|D_i = 0, p(X_i)]$ represents the outcome $\theta_{i,t}$ in countries that have not implemented IT, both conditional on the same probabilities $p(X_i)$ of IT adoption for a set of covariates X that have been estimated using the probit model mentioned earlier. With both terms on the right-hand side being observable, we are able to estimate the ATT in Eq. (4).

Figure A.9. in the online Appendix plots the estimated time-varying pass-through coefficients $\theta_{i,t}$, capturing the effect of interest rate changes on income inequality, measured by the Gini coefficient.[37] Table A.2. in the online Appendix reports the probit estimates (baseline with incremental controls for the exchange rate regime, government balance, and central bank independence). Across specifications, coefficients display the expected signs and are statistically significant.

Table 1 reports the ATT estimates from Eq. (4) using alternative matching algorithms. Columns (1) and (2) present nearest-neighbor matching with replacement ($n = 5$ and $n = 1$, respectively). Column (3) reports kernel matching, column (4) local linear matching, and column (5) radius matching ($r = 0.05$). Together, these estimates assess the effect of IT adoption on the pass-through measure. Indeed, the results show that adopting IT significantly increases the pass-through from interest rates to income inequality. Across matching methods and alternative specifications, estimated treatment effects range from 0.017 to 0.027, indicating a modest but robust amplification of the pass-through effect. The statistical significance of these estimates after addressing potential self-selection bias through PSM reinforces the argument that the choice of monetary policy framework matters for the distributional impact of monetary policy.

5. Conclusion

This paper contributes to the literature on the distributional effects of monetary policy by examining how the adoption of IT frameworks shapes the transmission of monetary policy to income inequality. By bridging the literature on the macroeconomic consequences of IT with recent work on the distributional impact of monetary policy, we provide robust evidence that the adoption of IT changes how monetary policy shocks affect income inequality. Our main finding is that expansionary monetary policy shocks reduce income inequality in countries operating under IT regimes, while in non-IT countries they generate a short-lived increase in income inequality that is economically meaningful at very short horizons but not persistent. This result is robust to controlling for key institutional features, including central bank transparency and independence, and holds across a wide range of robustness checks and alternative specifications. Addressing potential self-selection into IT regimes through propensity score matching further strengthens the credibility of this finding. In line with theoretical predictions, we also show that IT adoption is associated with more favorable macroeconomic outcomes, notably stronger output growth and improved price stability, compared to non-IT regimes. We

[37] As this is the first study to estimate these specific pass-through coefficients, no direct benchmarks exist. However, the estimated dynamics, volatile yet mean-reverting, are consistent with patterns documented in related pass-through literatures, such as exchange rate and oil price transmission to inflation (López-Villavicencio and Pourroy, 2019, 2022).

further show that the distributional effects of monetary policy operate through distinct transmission channels. Under IT regimes, the employment channel reinforces the transmission of expansionary monetary policy shocks toward declining inequality, while the financial channel exerts a comparatively weaker, dampening effect that partially offsets this decline.

A final caveat concerns the identification of monetary policy shocks. While the recursive ordering approach employed in this paper is standard in the VAR literature, it relies on assumptions about contemporaneous relationships and may not fully isolate exogenous shocks. Alternative identification strategies, such as high-frequency monetary policy surprises or narrative approaches (Nakamura and Steinsson, 2018; Jarociński and Karadi, 2020), would be desirable but are not feasible in our setting, as they require either high-frequency financial data or detailed country-specific historical records that are not consistently available for a broad cross-country panel, particularly at the annual frequency needed to match inequality data.[38] As a result, the use of VAR-based identification in this setting represents a compromise between coverage and precision. Our results should therefore be interpreted with some caution, especially for the non-IT group where heterogeneity is substantial. Reassuringly, the findings are robust to an alternative identification scheme based on sign restrictions. More broadly, these challenges are not unique to our setting. Recent work on financial shocks and inequality in the euro area (Ciganović et al., 2025) adopts similar identification strategies also constrained by the low frequency of distributional data and finds no meaningful differences between recursive and sign-restriction-based results. Future research could revisit these questions using richer data environments that allow for alternative identification strategies.

Declaration of competing interest

The authors declare that they have no known competing financial interests or personal relationships that could have appeared to influence the work reported in this paper.

Acknowledgment

We thank the Editor, Evi Pappa, and three anonymous referees for their insightful comments and suggestions.

Appendix A. Supplementary data

Supplementary material related to this article can be found online at https://doi.org/10.1016/j.euroecorev.2026.105359.

Data availability

Data will be made available on request.

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[38] See, for example, Amberg et al. (2022), who use Swedish administrative data combined with high-frequency monetary policy shocks. Such data requirements are not compatible with the cross-country setting of this paper.

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