Economic Quarterly—Volume 96, Number 3—Third Quarter 2010—Pages 229–258

Earned Income Tax Credit Recipients: Income, Marginal Tax Rates, Wealth, and Credit Constraints

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The Earned Income Tax Credit (EITC) has become the federal government's largest cash-assistance program for low-income families, making it the centerpiece of anti-poverty programs in the United States. Approximately 15 percent of households nationwide now qualify for the EITC (Hoffman and Seidman 2002). Moreover, unlike other government programs, the EITC is administered through the income tax filing process, which reduces any potential stigma associated with the program, and aids in ensuring high participation rates (Smeeding, Phillips, and O'Connor 2000). According to Eissa and Hoynes (2009), approximately $43 billion was allocated to 22 million families in the United States in 2007 through the federal EITC. This compares to $16.5 billion that was spent on more traditional welfare programs, such as Temporary Assistance for Needy Children (TANF).

The EITC is designed to augment income while encouraging work: The tax credit increases with earnings for low levels of household income. The size of the credit is such that, for low-income households that qualify, the EITC is a negative tax on earnings that often constitutes a significant portion of after-tax wage income. The EITC does appear to have been successful in both helping the working poor get out of poverty and encouraging work. Neumark and Wascher (2001), Ziliak (2006), and Simpson, Tiefenthaler, and Hyde (2009) provide evidence that the combined federal and state EITC helps

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families rise above the poverty line. In fact, the EITC has been estimated to have helped five million people out of poverty in 2005, including 2.6 million children.[1] Hotz and Scholz (2000) find that, compared to other poverty-reduction programs, the EITC is effective in raising the standard of living for low-income households, while keeping administrative costs relatively low.

However, the EITC phases out with earnings, until eventually a household no longer qualifies for it. The structure of the phase-out means that families earning more than $41,000 in 2008 will not qualify for the EITC, while all those earning less will. In addition, the credit targets families with children, and increases in generosity with the number of children in the household. For example, households with two or more children (in tax year 2008) earning $15,000 could qualify for up to $4,824 in federal earned income credits. In contrast, a childless single filer can receive only one-tenth of this amount, or at most $438. Thus, for those households with children and low earned income, the full refundability of the EITC ensures that it will represent a substantial addition to income.

In this article, we summarize the details of the EITC and describe the population of EITC recipients. Using Current Population Survey data, we estimate earnings and EITC benefits received by EITC recipients at various ages. Naturally, we find that because of the eligibility requirements, the earnings of EITC recipients are relatively similar across the age of recipients, which makes them differ systematically from non-recipients of the same age—whose earnings show a more pronounced “hump shape” with age. We then discuss how the EITC affects marginal taxes in the United States and summarize its theoretical and empirical effects on household labor supply decisions. Finally, we compare wealth levels of EITC recipients with non-recipients using data from the Survey of Consumer Finances (SCF), and find significant differences in their wealth distributions, with EITC recipients being substantially poorer. The fact that EITC recipients have relatively low wealth levels and low earnings relative to others in their age group suggests that they may be more likely to be borrowing-constrained than non-recipients. In fact, we find some evidence for this in our analysis of SCF data.

1. HISTORY OF THE EITC

In Table 1, we briefly summarize the history of EITC legislation. The EITC started as a modest program as part of the Tax Reduction Act of 1975.[2] The program was unique among tax credits as it was refundable so that poor families could utilize its benefits even if they owed little or no taxes. Unlike welfare programs such as Aid to Families with Dependent Children (AFDC), single

[1] Center on Budget and Policy Priorities: www.cbpp.org/cms/?fa=view&id=2505.

[2] For a more detailed history of the EITC, refer to Hotz and Scholz (2003).

Table 1 History of EITC Legislation

Year	Changes to the EITC
1975	Introduced temporary “work bonus” called the EITC
1978	Made EITC permanent
1986	General expansion (largest increase since its inception) and indexed for inflation; part of the Tax Reform Act
1990	General expansion by doubling the maximum credit and increased eligibility; added separate schedule for families with two or more children; part of OBRA
1993	General expansion (larger expansion for families with two or more children); added EITC for childless filers; part of OBRA
1997	Provisions made to improve compliance; part of Taxpayer Relief Act
2001	Changes to provide marriage penalty relief and promoted simplification; part of EGTRRA
2009	Expansion for families with three or more children and expanded eligibility for married couples; part of the American Recovery and Reinvestment Act

Sources: Hotz and Scholz (2003); Holt (2006); Tax Policy Center (2009).

parents as well as married couples were eligible for the program. The EITC went through minor changes in subsequent years, the most important being when it became a permanent provision of the Internal Revenue Code in 1978.

The Tax Reform Act of 1986 indexed the EITC to inflation and liberalized the EITC, helping, by some estimates, to remove over six million Americans from poverty (Ventry 2000). The Omnibus Reconciliation Act (OBRA) of 1990 increased the credit and added separate schedules for families with two or more children. The largest expansion of the EITC occurred in 1993, as part of the OBRA, in which the EITC was increased by an additional 25 percent. Families with two or more children experienced the largest increase in the credit, and childless filers could now qualify for the EITC. Both the size of the credit and the eligible population have grown over time, and were fueled by the passage of the Personal Responsibility and Work Opportunity Reconciliation Act of 1996, which replaced AFDC with Temporary Assistance for Needy Families (TANF). The United States experienced a 50 percent reduction in welfare rolls between 1993 and 2000, and Grogger (2004) finds that much of the drop is attributed to the EITC and reduction in welfare benefits.

Until 2001, the structure of the EITC was identical for single and married filers. However, as part of the Economic Growth and Tax Relief Reconciliation Act (EGTRRA) of 2001, married couples received larger benefits for larger ranges of income levels than single filers. The success of the federal EITC has led to the development of similar programs in 23 U.S. states and the District of Columbia, totaling an additional $2 billion (Levitis and Koulish 2008).

Table 2 EITC Parameters, Tax Year 2008

	Single, No Qualifying Children	Single, One Qualifying Child	Single, Two+ Qualifying Children	Married, No Qualifying Children	Married, One Qualifying Child	Married, Two+ Qualifying Children
Phase-In Rate	7.65%	34.00%	40.00%	7.65%	34.00%	40.00%
Phase-In Ends	$5,720	$8,580	$12,060	$5,720	$8,580	$12,060
Maximum Credit	$438	$2,917	$4,824	$438	$2,917	$4,824
Phase-Out Begins	$7,160	$15,740	$15,740	$10,160	$18,740	$18,740
Phase-Out Rate	7.65%	15.98%	21.06%	7.65%	15.98%	21.06%
Eligibility Ceiling	$12,880	$33,995	$38,646	$15,880	$36,995	$41,646

Source: Minnesota House Research Department.

Table 3 EITC Calculation by Phase

Phase	EITC
Phase-In	= Phase-In Rate * Income
Plateau	= Maximum Credit
Phase-Out	= Maximum Credit − Phase-Out Rate * (Income − Income Where Phase-Out Begins)

Finally, the American Recovery and Reinvestment Act of 2009 increased the credit for families with three or more children and expanded eligibility for married couples. Families making up to $48,250 in annual earnings can now qualify for the tax credit, with the maximum credit as high as $5,657 for a family with three or more children. This EITC expansion is expected to help an additional 650,000 households and 1.4 million children.[3]

2. STRUCTURE OF THE EITC

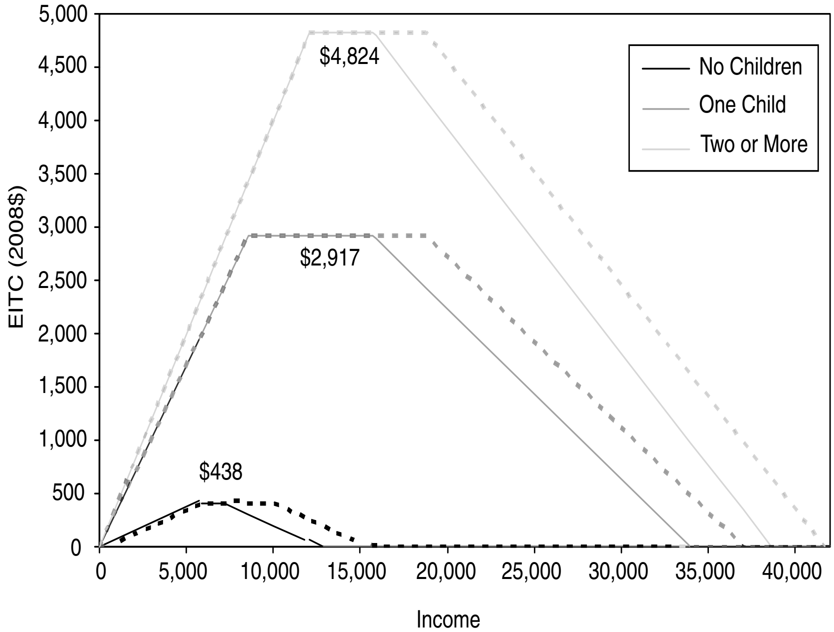
The EITC acts as an after-tax wage subsidy for low-income workers and depends on earned income, number of children, and marital status.[4] Earned income includes wages, salaries, tips, and other employee compensation; union long-term disability benefits received prior to minimum retirement age; and net earnings from self-employment. However, it does not include social security benefits, unemployment compensation, welfare benefits, scholarships, worker's compensation benefits, or pension/annuity income.

The EITC is structured in three phases: In the *phase-in* period, the credit increases with earnings; in the *plateau* period, the credit reaches a maximum and levels off; and in the *phase-out* period, the credit falls as the claimant's earnings rise. At the eligibility limit, the household earns no EITC. The EITC is separated into different levels for claimants with no children, those with one child, and those with two or more children. There are also different tax credits for different types of filers: Married couples filing jointly are eligible for slightly higher credit amounts in the phase-out period than single filers and have slightly larger income eligibility ranges. Table 2 presents the details of the EITC for tax year 2008 for different filing statuses (single or married) and number of children, and includes the maximum credits and earnings limitations. In Figure 1, we plot the amount of federal EITC that single and married households receive across various income levels: single filers are depicted by the solid lines, whereas married filers are depicted by

[3] Tax Policy Center (2009).

[4] Many of the poorest families are ineligible for the EITC since their earnings are too low to qualify and/or they do not have children (Hoffman and Seidman 2002).

Figure 1 EITC Structure by Income, Tax Year 2008



Note: Solid line represents single/head of household filers; dashed line represents married filers.

the dashed lines. To calculate the EITC in each phase, we use the equations in Table 3 along with the EITC parameters in Table 2.

As seen in Figure 1, the EITC significantly varies with the number of children present in a household. Childless filers receive less than one-eighth of the EITC than filers with one child and one-twelfth of filers with two or more children. The federal credit can represent up to 34 percent and 40 percent of income for filers with one and two or more children, respectively. In addition to the federal EITC, many states supplement, or match, with additional credits. As a result, if the taxpayer lives in a state that offers a state EITC, the total EITC (federal plus state) could be much larger; for example, New York residents receive an additional 30 percent of the federal credit. Also interesting is that the slope of the EITC function is steeper in the phase-in range than in the phase-out range. That is, an additional dollar of earned income rewards households in the phase-in range by giving them a credit, which can range from \$0.07 (for childless singles) to \$0.40 (for married couples with two children). In the phase-out range, an additional dollar of income results in a reduction in the credit, from \$0.07 (for childless singles) to \$0.21 (for married couples with two children).

The range of eligible income is much larger as the number of dependent children rises. As of 2008, married households with two children earning less than \$41,626 qualify for the EITC, compared to \$15,880 for childless couples. The maximum EITC does not vary with marital status, but the income eligibility ranges are slightly larger for married couples. In addition, the range of eligible income is much larger in the phase-out range so that more households are in the phase-out range than in the phase-in range. In fact, recent evidence suggests that married households are more likely to be in the phase-out range than singles, since they are more likely to have higher household income.

3. LABOR MARKET CHARACTERISTICS OF EITC RECIPIENTS

Using Current Population Survey (CPS) data from 2008, we analyze the labor market characteristics of EITC recipients and compare them to non-EITC recipients. We create household-level observations by matching individuals who are married to each other, and we restrict the sample to households where the household head is between the ages of 16 and 64 years. Households are classified into six different types, based on marital status (married or single) and number of children (no children, one child, two or more children). This classification is consistent with the structure of the EITC, as discussed in Section 2. We find that approximately 12.8 percent of households in our sample receive the EITC. Table 4a reports the mean annual wage and salary income, education level, and EITC amount for each household type, while Table 4b reports the fraction of each type in the sample. All of the means represent weighted averages using the household weights supplied by the CPS. It is important to note that 2008 CPS data corresponds to the 2007 tax year and that the CPS only reports estimated federal EITC and does not include any state EITCs.

Approximately 60 percent of EITC recipient households are single, with an equal distribution of single households having zero, one, and two or more children. This contrasts to married couple households, where the majority of EITC recipient households have two or more children. The amount of EITC varies significantly across household types. Single households with two children receive the most EITC (\$2,728), which constitutes the largest share of their annual income, at 15 percent. Households without children receive much less EITC, constituting only 6 percent of their annual income.

Table 4 EITC Recipients

4a: Labor Market Characteristics of EITC Recipient vs. Non-Recipient Households

	All	**Married, No Kids**	**Married, One Kid**	**Married, Two+ Kids**	**Single, No Kids**	**Single, One Kid**	**Single, Two+ Kids**
EITC Recipients:							
Mean Household Income	$15,194	$8,325	$18,700	$21,212	$7,024	$15,761	$17,421
St. Dev. of Household Income	$16,132	$8,100	$10,590	$11,225	$5,894	$9,739	$10,409
Percent of High School or Less	61.5%	70.5%	64.1%	68.5%	60.0%	54.1%	58.3%
Percent with Two Earners	26.3%	9.6%	24.6%	30.1%	n/a	n/a	n/a
Average EITC	$1,782	$495	$1,812	$2,623	$423	$1,808	$2,728
EITC as Percent of Income	11.7%	5.9%	9.7%	12.4%	6.0%	11.5%	15.7%
Non-EITC Recipients:							
Mean Household Income	$47,235	$68,549	$83,372	$94,271	$23,696	$32,125	$31,723
St. Dev. of Household Income	$49,653	$67,884	$71,052	$79,822	$32,305	$47,998	$51,997
Percent of High School or Less	39.7%	36.1%	34.2%	29.6%	43.4%	46.6%	48.3%
Percent with Two Earners	65.1%	56.8%	70.5%	71.3%	n/a	n/a	n/a

4b: Distribution of Households in the CPS

	Married, No Kids	**Married, One Kid**	**Married, Two+ Kids**	**Single, No Kids**	**Single, One Kid**	**Single, Two+ Kids**	**Sum**
EITC Recipients:							
Percent of All Households	0.59%	1.21%	3.12%	3.15%	2.32%	2.45%	12.83%
Percent of EITC Recipients	3.9%	9.6%	26.8%	19.8%	19.4%	20.6%	100.00%
Non-EITC Recipients							
Percent of All Households	14.60%	7.91%	12.40%	46.26%	3.81%	2.19%	87.17%
Percent of Non-EITC Recipients	14.77%	10.57%	18.61%	48.16%	4.91%	2.99%	100.00%

Notes: Household data constructed using 2008 CPS; 16–64-year-olds, 2008 dollars. Means are weighted using the household weight “hhwt.”

Much of the variation in the EITC across household types is because of differences in annual income. Not surprisingly, married households earn more than single households since there is the potential for two earners. It is interesting to note, however, that the share of married households that have two earners is quite low for EITC recipients, compared to non-recipients. For example, approximately 30 percent of married households with two children who receive the EITC have two earners, while 71 percent of non-recipients have two earners. This could be due to the fact that the majority of two-earner households surpass the income qualifications of the EITC. Or, it could be that EITC-recipient households choose not to have a second income since they receive the EITC.

Another interesting feature is that household earnings for EITC recipients increase with the number of children, and this occurs for both married couple households and single parent households. The difference in annual income between childless households and households with children is much larger for EITC-recipient households than for non-recipient households.

Even though single households that receive the EITC earn less than married households, they tend to be more educated (for married households, we use the education level of the household head). Approximately 10 percent fewer single households have a high school degree or less compared to married households and this is independent of the number of children. This is not the case for non-recipient households: Single households that do not receive the EITC are more likely to only have a high school education than married households.

Thus, the EITC likely has the largest impact on households with children since the EITC is much larger for these households as a share of their annual income and more than 75 percent of EITC recipient households have children. Single households represent the majority (60 percent) of EITC recipient households, and tend to be more educated than married EITC households, which contrasts with the general population. EITC recipient households are much less likely to have two earners than non-recipient households.

4. EITC AND INCOME BY AGE

We now analyze how the EITC changes across recipients of different ages. Since the EITC targets low-income families, it will disproportionately affect younger households of child-rearing age. However, households may qualify for the EITC at any stage of their life, as long as they have earned income that is below the income limit. Importantly, there is no limit to the amount of benefits received over a lifetime nor is there a time limit.

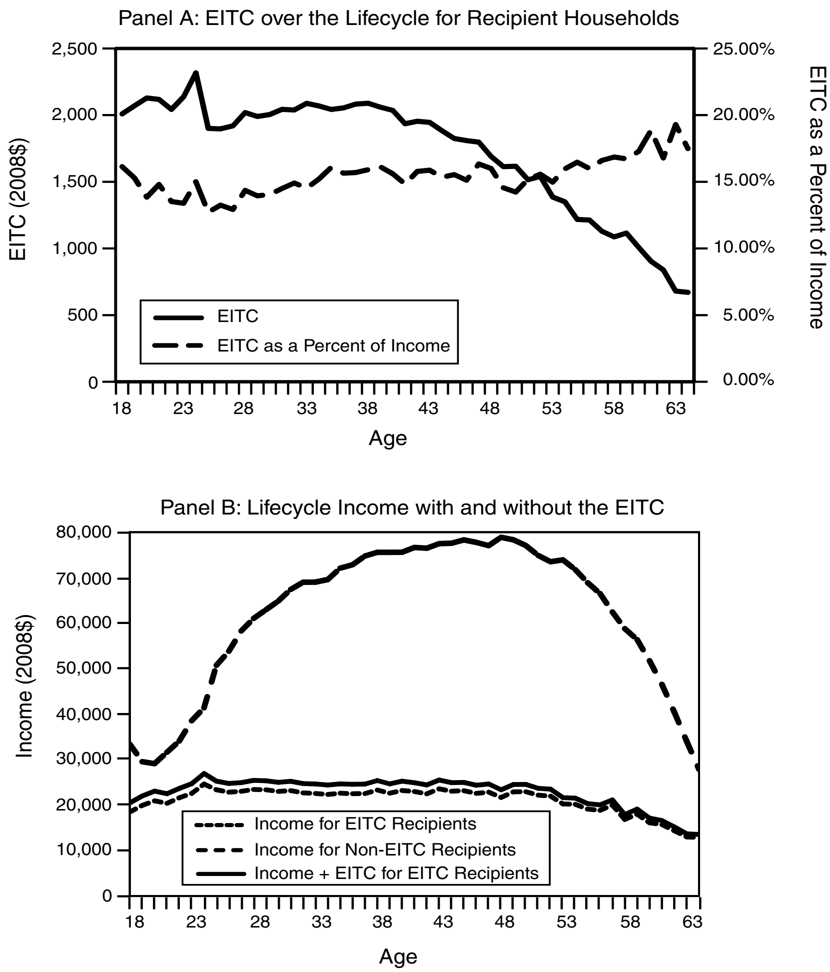
We analyze the pool of EITC recipients between 1992–2008 and catalog how the EITC varies across households of different ages in a shortened panel. Specifically, we estimate the average income/EITC (in 2008 dollars) for each household at each age in each year of the CPS (using the household weights supplied by the CPS). Then, we calculate the average income/EITC across the panel by age; to do this, we account for the distribution of households at each age across the panel. This yields an estimate for income/EITC, conditional on receiving EITC, at each stage in the lifecycle for the typical household in the CPS.

While the preceding is useful, it is an imperfect measure of the effect of EITC on lifetime earnings. It abstracts from any cyclical effects that individuals experience in earnings (such as business cycles, changes in skill premium, or occupational transitions) that occurred prior to 1992 for older cohorts (for example, changes in earnings profiles for individuals born before 1974 are not accounted for prior to 1992). However, our method accurately accounts for the drastic changes that occurred in the EITC during this period. In addition, our estimates provide a sense of how the EITC changes by age and what households can expect as they age, should they qualify at later dates.

In Figure 2a, we plot the average EITC for households that receive the EITC at each age between 18–64 using 1992–2008 CPS data (the age of the household head is used); we also plot the EITC as a percent of earnings (labor earnings and EITC) in the same figure. A few interesting findings emerge. The EITC is high for households headed by very young adults (age 18–25), relatively constant for households in their thirties (at approximately $2,000 in 2008 dollars), and then declines precipitously as we look at households in their late thirties and beyond. By the time households are in their fifties and sixties, the average amount of EITC is just over $500. Thus, the amount of EITC that households receive declines over the course of their lifetimes. However, the interaction of the qualification requirements and the structure of benefits ensure that the EITC remains a relatively constant fraction of recipients' earnings, at approximately 15 percent, for most of their lives. While the typical EITC transfer is largest for the youngest recipients in our sample, the EITC represents a significant fraction of annual earnings (at least 15 percent) throughout most of a recipient's working life. In addition, the EITC represents an even larger proportion of the income of older EITC recipient households. For example, for EITC recipients in their late fifties, the EITC increases as a percent of earnings to approximately 18 percent. This is likely due to the fact that households that qualify for the EITC at this age have very low incomes since they likely face the income thresholds applicable to those with no children.

The patterns in EITC receipt across different age groups arise from two factors: child-rearing stages and fluctuations in income over the lifetime. A typical lifetime earnings profile exhibits a hump shape, where earnings are low early in life, increase dramatically through the twenties and thirties, level off through the forties, and start to decline in the fifties and sixties. This is exactly what we observe for non-EITC recipient households in the CPS sample. In Figure 2b, we plot household earnings (wages and salary) profiles

Figure 2 EITC Recipients and Non-Recipients Across Ages



Notes: Household data constructed using 1992–2008 CPS; 16–64-year-olds, 2008 dollars. Means are weighted using the CPS household weight "hhwt."

for non-EITC recipients and EITC recipients. By construction of the eligibility requirements for EITC, however, those receiving it at various ages are much more similar to each other than are non-recipients of differing ages. Amongst recipients, the highest levels of benefits accrue to the young, typically around age 25. Older recipients generally earn smaller amounts, primarily as the number of dependents they may claim falls.

5. MARGINAL INCOME TAX RATES

The EITC represents a negative income tax for households that qualify for it. Thus, for low income levels, marginal income tax rates are negative. Using data from TAXSIM version 9.0 from the National Bureau of Economic Research,[5] we calculate the marginal income tax rates for all single and married households with no children, one child, and two children (i.e., dependents exemptions) for tax year 2008.[6] The marginal income tax rate is for adjusted gross income only and does not include Federal Insurance Contributions Act (FICA) contributions (i.e., Social Security and Medicaid).

In Figure 3, we plot the marginal tax rates across income levels for single and married filing status earning up to $100,000 and differentiate households based on the number of children they claim as dependents. As you can see in the first panel for married households with two or more children, for low levels of income, the marginal tax rate is −40 percent for both single and married filers, which represents the phase-in rate for the EITC. As incomes reach $13,000, the marginal rate is 0 percent (in the plateau region). For households with income above $13,000, the marginal tax rate becomes positive and gets quite large quickly. For married households with incomes between approximately $19,000–$25,000, the marginal tax rate jumps to 21 percent, which represents the EITC phase-out rate. That is, at the margin, these households are experiencing a 21 percent reduction in their EITC for any additional income they earn in this range. For married households with incomes between approximately $25,000–$40,000, the marginal income tax rate increases to 31 percent, which represents the EITC phase-out rate plus the lowest income tax bracket of 10 percent. For married households with two children earning $41,000, they face the phase-out rate and the next highest tax bracket of 15 percent, making their marginal tax rate 36 percent. Thus, the phasing out of the EITC leads to dramatic increases in the marginal income tax rates for these households. For married households above $41,000, they no longer qualify for the EITC; hence, they face significant reduction in their marginal tax rates, at 15 percent (in the second income tax bracket). As household income approaches $90,000, the marginal tax rate increases to 25 percent for married filers.[7,8] Single taxpayers with two children experience similar jumps in the marginal income tax rates, but for lower levels of income than married households.

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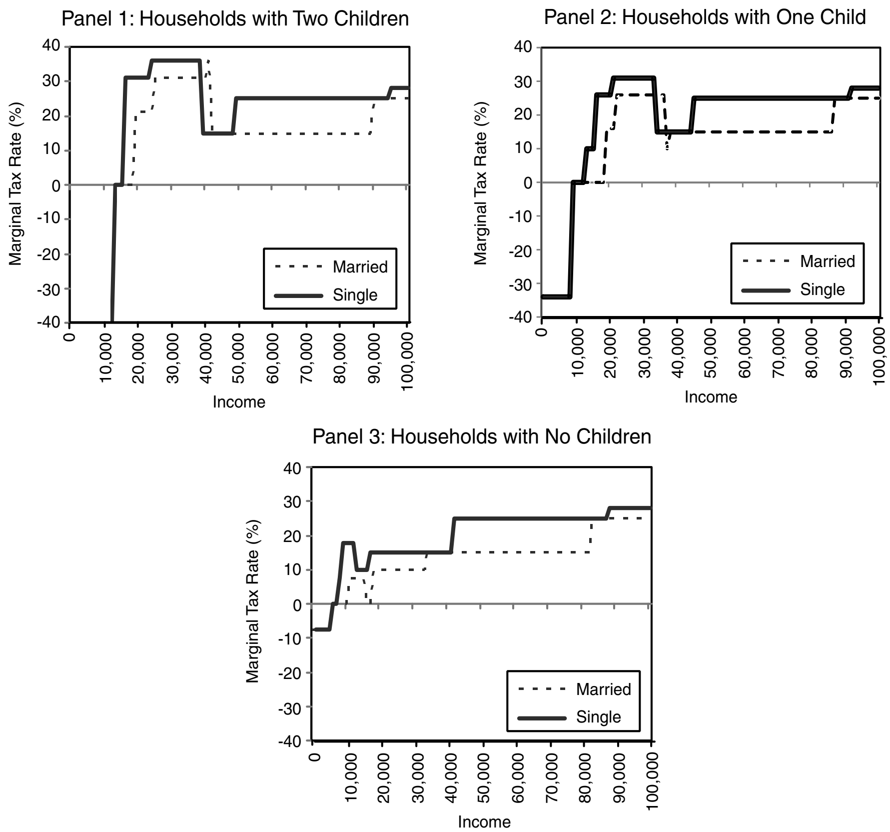
5 www.nber.org/˜taxsim/taxsim-calc9/index.html.

6 We follow the methodology of Hotz and Scholz (2003), Romich (2006), and Eissa and Hoynes (2009) in generating the marginal tax rate schedule.

7 Marginal tax rates in the United States increase up to 35 percent for household incomes up to $357,000 (in 2008). However, we focus on income tax rates for low- and middle-income households.

8 If we were to include FICA contributions, the entire marginal tax curve would shift upward by 7.65 percentage points across all income levels.

Figure 3 Marginal Income Tax Rates



Source: TAXSIM 9.0, 2008 tax year.

The second panel in Figure 3 shows the marginal income tax schedule for married and single households with one child. The figure is similar for those with two or more children, however, the marginal rates are slightly lower across all income levels. For example, the poorest households with one child face a marginal tax rate of −34 percent (compared to 40 percent for households with two or more children). In addition, marginal tax rates for those earning between $20,000–$40,000 are approximately 5 percentage points lower for those with one child, because of differences in the slope of the phase-out rate (the phase-out rate is steeper for those with more children, as documented in Table 2). As households go beyond EITC eligibility, the marginal income tax schedule does not vary with the number of children. Once again, these households experience significant reductions in their marginal tax rates as soon as they are ineligible for the EITC.

In the last panel of Figure 3, the income tax schedule is quite different for those with no children compared to those with children. Recall that the EITC is much less generous for childless households. Thus, the negative marginal rates are quite low (in absolute value terms) for the poorest households. Also notice that the increases in the marginal rates are not as extreme for childless singles; as a result, these households do not experience significant reductions in their marginal tax rates as they become ineligible for the EITC (for incomes above $15,800 for married households). Beyond EITC eligibility, they face the same marginal income tax rates as households with children.

Our analysis of the marginal income tax schedule for EITC recipients uncovers a few interesting points. First, the very poorest households with children (those earning below $12,000) experience large negative income tax rates (in absolute value terms) because of the EITC. Second, single parent households that receive the EITC face some of the highest (positive) marginal income tax rates in the United States (Ellwood and Liebman 2000); for example, a single mother with two children earning $35,000 pays a marginal income tax rate of 36 percent (in 2008). These high marginal tax rates can be attributed to the phasing out of the EITC and the progressive income tax schedule (Romich 2006). Married households with children face slightly lower marginal tax rates than single households with children. Third, once households with children no longer qualify for the EITC, their marginal income tax rates drop significantly, and once they surpass EITC eligibility, marginal income tax rates no longer depend on the number of children in the household.

6. LABOR SUPPLY RESPONSE TO EITC

As a wage subsidy, the EITC has the potential to affect both the decision to work (i.e., the extensive margin) and the number of hours worked (the intensive margin). In a static labor-leisure model, the EITC increases the marginal value of working (i.e., the after-tax wage rate). Thus, in theory, the EITC will increase labor market participation because of the substitution of work for leisure. However, the effects of the EITC on hours worked are theoretically ambiguous. We follow the formulation in Eissa and Hoynes (2009) in extending the labor-leisure model to include the EITC.

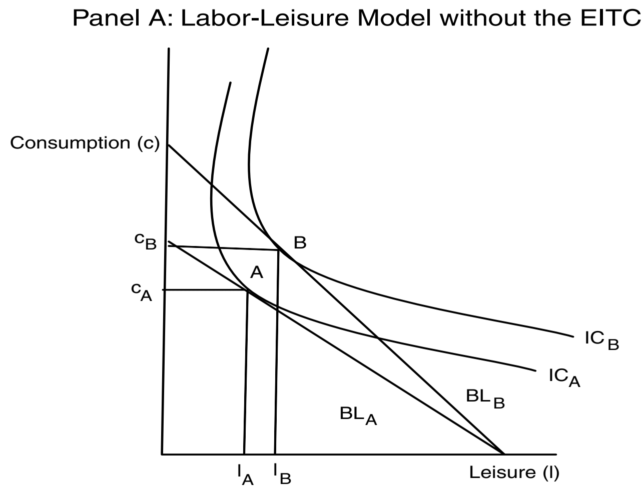
Consider a representative household within the traditional labor-leisure model, where the household unit decides how much to work. The household could constitute one or more workers, where the tradeoff to working is household leisure. The budget constraint (without the EITC) is depicted by: $c = \tilde{w} * n$, where c represents consumption, $\tilde{w}$ represents after-tax wages, and n represents labor hours. Households have T units of time to devote to labor (n) and leisure (l); $T = n + l$. The slope of the budget constraint, and hence the cost of pursuing an additional unit of leisure, is $\tilde{w}$ units of consumption. In Figure 4a, we plot the budget constraint with leisure on the x-axis and consumption on the y-axis (BL_A). Plotting an indifference curve on this graph (with all of the standard assumptions for utility) provides the equilibrium quantity of leisure (l_A) and consumption (c_A), at point A. If after-tax wages rise because of a reduction in the marginal income tax rate, the budget line gets steeper (rotates to BL_B). For the same household, the equilibrium quantity of leisure/labor may rise or fall because of the tax cut. The substitution effect reduces leisure, and hence raises labor supply. The income effect raises leisure and lowers labor. The net effect depends on the relative size of each effect. In the diagram, the income effect dominates such that labor supply falls (leisure increases) in response to a tax cut ($l_B > l_A$).

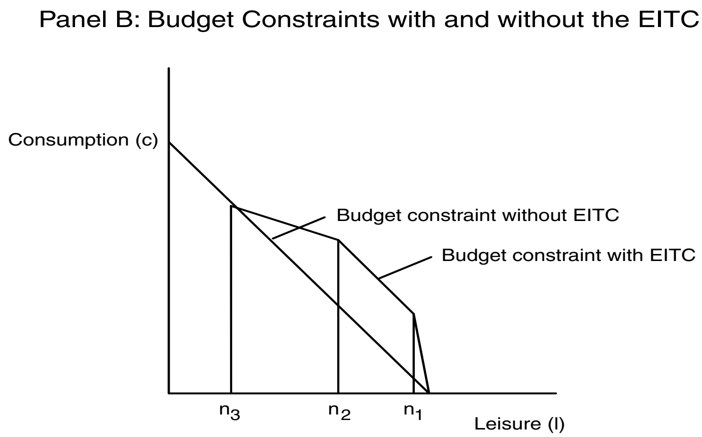
The EITC changes the after-tax wage rate ($\tilde{w}$) for different levels of leisure/labor. For low levels of labor, when the household receives a tax credit (i.e., a negative tax) for each additional unit of labor, the after-tax wage is $\tilde{w} = w\,(1 + t_s)$, where $t_s > 0$ is the phase-in rate. For higher levels of labor in the plateau region, the after-tax wage is simply w since the EITC is constant in this range; that is, $\tilde{w} = w$ where households receive a transfer, Tr. During the phase-out region, the after-tax wage is $\tilde{w} = w\left(1 - t_p\right)$; the EITC falls for each additional unit of labor at the rate $t_p > 0$. For very high levels of labor, the after-tax wage returns to w once again. Thus, the budget constraint is as follows: $c = w\left(1 + \tau_p\right) * n$ for $n \in (0, n_1)$; $c = w * n + Tr$ for $n \in [n_1, n_2)$; $c = w\,(1 - t_s) * n$ for $n \in [n_2, n_3)$; $c = w * n$ for $n \in [n_3, T)$; where Tr is the maximum EITC and n_i represents different quantities of labor. The EITC budget constraint, as plotted in Figure 4b, is kinked at each quantity of labor n_i in which $\tilde{w}$ changes.

By comparing the budget constraint with and without the EITC in the various ranges of labor supply, we can determine the theoretical effects of the EITC on hours worked. First notice that for households that do not work ($l = T$), the EITC is 0 and has no effect on the household's budget constraint. However, for those households that choose to work very little (i.e., $n = \varepsilon$, where $\varepsilon \in (0, n_1)$), the slope of the budget line gets steeper. Here, there is a positive substitution effect and no income effect. Thus, the EITC may influence some households to enter the labor force, leading to a positive effect on the extensive margin.

However, the effects of the EITC on the intensive margin are more complicated. In the phase-in range, the slope of the budget constraint is higher with the EITC ($\tilde{w} > w$ since $t_s > 0$); thus, a negative income effect and a positive substitution effect are both at play, making the effects on hours worked ambiguous. Those in the plateau region receive the same amount of credit if they earn more income, and hence a pure income effect occurs in which higher income reduces the incentive to work. In the phase-out range, the slope of the budget constraint is flatter than without the EITC ($\tilde{w} < w$ since $t_p > 0$).

Figure 4 EITC and Labor Supply





Here, a negative substitution effect influences households to substitute leisure for hours worked. In addition, a negative income effect may reduce hours worked even more. Thus, households in the phase-out region unambiguously reduce hours worked. Since a majority of EITC recipient households fall in the flat or phase-out region, it is likely that the overall effects of the EITC on hours worked are negative (Hotz and Scholz 2003). For those with income beyond the phase-out region ($n \in [n_3, T)$), their return to an additional hour of work is w, so that some of them may choose to restrict labor hours to be eligible for the EITC, once again leading to a negative extensive margin effect.

Of course, the magnitude of these responses depends on the elasticities of labor supply. High elasticities lead to larger labor supply responses, and labor supply elasticities vary across different types of people. For example, the uncompensated elasticity of labor supply is higher for women than for men and the elasticity on labor force participation is larger than the elasticity of hours (Evers, Mooij, and Van Vuuren 2008). Thus, the quantitative effects of the EITC on both the extensive and intensive margins of labor supply decisions depend critically on the presumed elasticities of labor supply.

There is a large empirical literature that examines the effects of the EITC on labor supply, with most of the work focusing on single mothers. For a more detailed summary of this literature, refer to Holt (2006) and Hotz and Scholz (2003). The evidence indicates that the EITC does in fact increase labor force participation, especially for single mothers (Meyer 2001), leading to positive effects on the extensive margin. In fact, the EITC has led to a dramatic increase in employment rates for single mothers during the 1980s and 1990s (Eissa and Leibman 1996; Meyer 2001; Grogger 2004). However, the effects of the EITC on the intensive margin are less clear in the data, with most studies not finding a significant change in hours worked because of the EITC. The most relevant work here is that of Cancian and Levinson (2005), who study a natural experiment arising from the fact that one U.S. state (Wisconsin) altered the generosity of its matching of the federal EITC. They argue that there is essentially zero effect on hours. There is some evidence, however, suggesting that single mothers may work more in response to the EITC since they are likely to be in the phase-in region where marginal income tax rates are negative (Eissa and Liebman 1996). Married women, however, who typically fall in the phase-out range, may work fewer hours as a result of the EITC rates (Ellwood 2000; Eissa and Hoynes 2004).

Very few studies analyze the labor market effects of the EITC on married couples; notable exceptions include Eissa and Hoynes (2004, 2009). They find that the EITC has small negative effects on both the extensive and intensive margins for married couples. However, the EITC has differential effects on primary and secondary earners. For example, increases in the EITC lower both the participation rates and hours worked for secondary earners since these households are usually being phased out of the EITC, where the returns to working more are relatively low.

There seems to be some consensus in the empirical literature that the EITC has positive effects on the extensive margin for households and little to no effect on the intensive margin. Studies have shown that the labor supply of low-income households is generally unresponsive to high marginal tax rates (Keane and Moffitt 1998; Gruber and Saez 2002); this compares to high-income workers who are quite responsive to tax rates. Perhaps low-income workers cannot adjust their work hours because of their job structure (Romich 2006). Or perhaps these workers do not realize the high marginal tax rates because of the complexity of the income tax and benefits structure in the United States. Recent theoretical work in a separate but related context suggests that a central force may be that low-income households are typically *low-wealth* households. As a result, these households will often be close to a borrowing constraint. Consumption theory predicts that such households will work in a manner insensitive to current wages, as the value of lowering the likelihood of a binding borrowing limit (by working and reducing consumption) will be high. The work of Pijoan-Mas (2006) suggests that this may be exactly the case, as he is able to rationalize a relatively high willingness of households to substitute labor intertemporally, with a low aggregate correlation between wages and hours. In ongoing work, Athreya, Reilly, and Simpson (2010) utilize this insight and embed households into a setting in which they face uninsurable risks and liquidity constraints, and find that, indeed, the disincentives to labor supply arising from the EITC are not strong.

7. WEALTH DISTRIBUTION OF EITC RECIPIENTS

As documented above, EITC recipients earn much less over their lifetimes than the general population. This will have important effects on their wealth holdings. In addition, their wealth level may affect their labor supply decision, as discussed above. In this section, we use the 2007 SCF to compare the distribution of wealth for EITC recipients and non-recipients, and then analyze differences across the six different types of households. Wealth is defined as household net worth, which is the difference between total assets and total debt.[9] The SCF does not report anything related to the EITC. However, we calculate the imputed EITC level that households would have received in tax year 2006 using the household structure and wage/salary income reported by the SCF. That is, we feed the parameters of the federal EITC program into the SCF to generate a proxy for the amount of EITC each household is eligible to receive. However, it should be made clear that we cannot observe directly if each household received the EITC—we know only whether or not they qualified for the EITC and, if they qualified, how much EITC they should have received.

All of the usual caveats apply when using the SCF data, in that it is a small sample and is not representative of the U.S. population at large. Our sample of the 2007 SCF contains 3,458 households compared to 86,259 households in the 2008 CPS (recall that we restrict the analysis to household heads between 16 and 64 years old and use the individual-level data in the CPS to create household-level observations). It is well-known that the SCF oversamples wealthy and married households. For example, when comparing the

[9] We use the SCF definition of net worth, as used in various Federal Reserve Bulletin articles, including Bucks et al. (2009).

distribution of household types between the CPS (reported in Table 4b) and the SCF (in Table 5b), it is evident that married households are oversampled in the SCF compared to the CPS and that single households are undersampled (and especially childless singles and single parents with one child). Surprisingly, the SCF just slightly oversamples households that are eligible for the EITC; they represent 12.8 percent of the CPS sample and 16.4 percent of the SCF sample. Also, the SCF does surprising well in capturing an accurate distribution of EITC recipients across household types and their mean income and EITC levels, compared to the CPS. This provides support to our use of the SCF to analyze EITC recipients. All of the reported means are reported in 2007 dollars and are weighted using the replicate weights produced by the SCF.[10]

In Table 5a, we report mean net worth (i.e., wealth), assets, debt, and income across household types. Not surprisingly, households that qualify for the EITC have much less net worth, assets, and debt than non-recipient households, and the difference is astounding. Mean net worth of EITC recipients is \$103,753 (in 2007 dollars) compared to \$580,245 for non-recipients. Some of the difference in net worth between EITC and non-EITC recipients can be explained by differences in income and age: EITC recipients earn 23 percent of what non-recipients earn, on average, and are almost six years younger. Somewhat interesting is that mean debt level for EITC recipients is \$45,755, which represents 2.6 times their annual salary, compared to non-recipients whose debt-to-income ratio is approximately 1.7. Thus, debt-to-income ratios are quite high for households that qualify for the EITC.

In Table 6, we report mean wealth by quartiles for both EITC and non-EITC recipients. First, notice that households in the lowest quartile of EITC recipients have average negative wealth of −\$16,617. In fact, 18.4 percent of households in the EITC sample have negative net worth. However, there is a significant amount of heterogeneity in the first quartile, as evidenced by the large standard deviation. This compares to the lowest quartile of non-EITC recipients, whose mean wealth level is \$1,899 and standard deviation is \$324. Second, notice that the wealth distribution for EITC recipients is much tighter than for non-recipients. The ranges of wealth in each quartile are much smaller and the standard deviations are generally lower (with the exception of the first quartile of EITC recipients). Third, the majority of EITC recipients hold very little wealth; those in the third quartile of wealth hold on average only \$24,038 in net worth, compared to non-recipients in the third quartile who hold more than \$250,000. Only the top quartile of EITC recipients has a significant amount of wealth. In fact, only 20.3 percent of EITC recipients

[10] For a full discussion of the importance of weights in the SCF, refer to Kennickell (1999).

Table 5 Balance Sheets of EITC Recipients and Non-Recipients

5a: Assets, Debt, and Net Worth of EITC Recipient vs. Non-Recipient Households

	All	Married, No Kids	Married, One Kid	Married, Two+ Kids	Single, No Kids	Single, One Kid	Single, Two+ Kids
EITC Recipients:							
Mean Net Worth	$103,753	$284,403	$204,918	$118,468	$67,574	$56,102	$49,837
Mean Assets	$149,507	$359,963	$255,239	$179,050	$86,545	$89,365	$96,465
Mean Debt	$45,755	$75,560	$50,321	$60,582	$18,971	$33,263	$46,628
Mean Household Income	$17,593	$6,199	$21,818	$22,502	$6,990	$18,903	$19,070
Mean (Imputed) EITC	$1,778	$231	$1,440	$2,409	$277	$1,720	$2,726
Mean Age	38.5	46.6	37.2	37.5	37.1	40.4	38.1
Non-EITC Recipients:							
Mean Net Worth	$580,245	$803,447	$621,345	$737,654	$275,437	$351,416	$223,309
Mean Assets	$708,564	$929,270	$790,176	$933,762	$334,930	$448,206	$296,280
Mean Debt	$128,319	$125,823	$168,830	$196,108	$59,493	$96,790	$72,971
Mean Household Income	$76,686	$87,916	$95,962	$105,640	$38,071	$50,373	$35,849
Mean Age	44.3	46.9	43.6	41.3	44.8	47.2	41.5

5b: Distribution of Households in the SCF

	Married, No Kids	Married, One Kid	Married, Two+ Kids	Single, No Kids	Single, One Kid	Single, Two+ Kids	Sum
EITC Recipients:							
Percent of All Households	0.97%	2.20%	4.88%	2.53%	2.46%	3.38%	16.42%
Percent of EITC Recipients	5.9%	13.4%	29.7%	15.4%	15.0%	20.6%	100.00%
Non-EITC Recipients							
Percent of All Households	25.47%	12.66%	22.68%	17.42%	2.83%	2.52%	83.58%
Percent of Non-EITC Recipients	30.47%	15.15%	27.13%	20.84%	3.39%	3.02%	100.00%

Source: Authors' calculations using the 2007 SCF. Means are weighted, in 2007 $.

hold more than the average wealth level for EITC recipients ($103,753). This compares to non-recipients, where 41 percent hold more than the average wealth level of $580,245 and 69 percent have more wealth than the average EITC recipient.

There is significant variation in wealth across household types, as illustrated in Table 5a. Married households have three times as much wealth as single households, with the largest difference for households with no children. It is likely that most of the wealth held by married households with no children is comprised of housing wealth since this group is relatively old. In addition, mean household wealth is smaller for households with more children despite higher earnings, and this effect is particularly large for married households. Thus, mean wealth levels for single households are quite low but are not that different for those with and without children. For married households, households with children have higher earnings but significantly less wealth compared to those without children. This is partially explained by age differences across married households—those without children are approximately nine years older than those with children. In addition, single households without children earn the least income of any group, but are not the poorest type of household in terms of net worth. Single households with two or more children have the lowest net worth in both the EITC and non-recipient samples.

Our analysis documents several interesting findings about the wealth holdings of EITC recipients. Not surprisingly, we find that EITC recipients hold very little wealth: EITC recipients, on average, hold only one-fifth of the wealth of non-EITC recipients. In fact, the bottom quartile of EITC recipients hold negative wealth on average, while the bottom quartile of non-recipient households have small, positive wealth holdings. However, debt-to-income ratios of EITC households are significantly higher than those of non-recipients (2.6 compared to 1.7 on average). We find that married households that are eligible for the EITC hold more wealth than single households, and wealth holdings decrease with the number of children in the household.

8. EITC AND CREDIT CONSTRAINTS

Based on the data presented in Figure 2b, the EITC increases earnings for recipients during every year of their working life and more so in early life. In a typical lifecycle model of savings and consumption, a household would save in periods when income is high, and borrow when income is low. As a result, the EITC allows low-income families to smooth consumption over their lifetimes. At higher frequencies, such as within a given year, the EITC can help, even though most families receive the EITC in lump sum when they

Table 6 Wealth Distributions: EITC Recipients vs. Non-Recipients

	Mean	**St. Dev.**	**Min.**	**Max.**	**Mean Income**	**Mean Age**
EITC Recipients:						
Bottom Quartile	−$16,617	$1,860	−$473,700	$170	$14,938	34.0
Lower Middle Quartile	$3,489	$85	$190	$7,560	$15,919	33.7
Upper Middle Quartile	$24,038	$531	$7,630	$51,400	$20,507	38.6
Upper Quartile	$404,272	$24,215	$52,120	$615,000,000	$19,014	47.7
Non-EITC Recipients:						
Bottom Quartile	$1,899	$324	−$118,999	$24,120	$34,055	37.9
Lower Middle Quartile	$75,329	$697	$24,130	$141,500	$51,829	42.6
Upper Middle Quartile	$253,637	$1,467	$141,520	$396,210	$76,599	46.5
Upper Quartile	$1,991,197	$33,646	$396,300	$806,000,000	$144,308	50.4

Source: Authors' calculations using the 2007 SCF. Means are weighted, in 2007 dollars.

file their tax returns.[11] In addition, households may borrow against their EITC, knowing that they will be receiving it later. Alternatively, households may save their EITC for future consumption.

The ability of households to smooth (bring forward) an expected EITC lump-sum payment that is made at the time of one's annual income tax payment depends on the household's ability to borrow. For those who can borrow, the EITC may act as insurance against income, employment, or health shocks, for example. If, on the other hand, households face significant borrowing constraints, they may not be able to borrow against their EITC, and so, while the EITC still provides low frequency smoothing, it may not assist consumption smoothing efforts within shorter periods, for example one calendar year.

Direct evidence on the extent to which EITC recipients are credit constrained is not possible, given current data limitations. Moreover, credit constraints are generally very difficult to identify. Typically, the measurement of credit constraints in any given study relies on a particular theory of consumption to identify consumption or savings movements that appear "anomalous," such as the large "excess sensitivity" literature on the 1980s for the path of aggregate consumption (see Deaton 1992). A handful of articles find evidence that suggests that those who share demographic characteristics with the EITC recipients are likely to be credit constrained. For example, the results of Jappelli (1990) indicate that lower income, wealth, and age are all associated with higher likelihoods of being credit constrained, all key features of the EITC population as documented above. Souleles (1999) finds that households that receive tax refunds and are liquidity constrained experience significant increases in nondurable consumption at the time of refund receipt. Barrow and McGranahan (2000) discover a seasonality of consumption behavior that is consistent with the timing of the receipt of the EITC, especially for durable goods. Berube et al. (2002) discuss the proliferation of paid tax preparation services and refund loans (at relatively high interest rates) for EITC recipients, suggesting that these households lack financial services and, hence, access to credit. Finally, Elliehausen (2005) analyzes survey data from households that use refund anticipation loans (RALs). He finds that EITC recipients who use RALs are less likely to use various types of credit (including car loans, bank and retail credit cards, and mortgages) than other RAL households. In addition, Elliehausen (2005, 52) reports that:

> Nearly half of EITC recipients that obtained RALs reported being turned down or limited by a lender in the last five years, and a little more than half said that they had thought about applying for credit but did not because they thought that they would have been turned down. These

[11] The advance EITC allows them to receive their EITC throughout the year in their paycheck, but very few households participate in this option (Romich and Weisner 2000).

> percentages are more than two times the percentage of all households experiencing turndowns or limitations and more than three times the percentage of all households perceiving limitations in credit availability.

However, no study has provided direct evidence of the extent to which EITC recipients are credit constrained.

Using 2007 SCF data and following Jappelli (1990), we use a set of questions from the SCF that provide a sense of the severity of credit constraints that EITC recipient households face. We use the following four measures:

1. *Bad credit*: For households that do not have a checking account, the SCF asked why. If the response was because of credit problems, bankruptcy, and/or does not meet qualifications for an account, then a value of 1 was assigned.

2. *Credit card balances*: This is the total value of credit card balances held by households. Credit card balances consist of the amount outstanding on all credit cards and revolving store accounts after the last payment. Balances do not include purchases made since the last account statement.

3. *Late payment for 60+ days*: This was assigned a value of 1 if the household had any debt payments more than 60 days past due in the last year.

4. *Has no checking account*: This was assigned a value of 1 if the household did not have a checking account.

Certainly, these four measures are not perfect predictors of being credit constrained. For example, some households choose not to have a checking account for reasons that are unrelated to their credit status. However, not having a checking account will undoubtedly lead them to have less access to credit in the future; without a checking account, many banks are not willing to issue personal loans and/or mortgages. That is, the causality between these measures and the likelihood of being credit constrained is unclear; however, if we find some correlation between these measures and the EITC, it may shed some light on the extent to which EITC households are or will be able to borrow. Similarly, credit card balances are an imperfect measure of credit constraints; lower balances may imply less willingness to use credit cards and/or acquire debt, and not less ability to borrow. But it may also indicate that they have lower credit limits, suggesting tighter borrowing constraints. Of the four measures above, having bad credit and late payments are perhaps the most accurate measures of credit constraints since both will lead to lower credit scores and, hence, worse credit terms.

In the analysis that follows, we compare these four measures for households that receive the EITC versus non-recipient households. As we document in Section 1, EITC-recipient households are younger, less educated, and have more children than non-recipient households; as a result, they are poorer. Obviously, having fewer current and, especially, future resources to borrow against will make it more difficult for EITC-recipient households to borrow. Nonetheless, it is useful to know the extent to which any household is likely to be constrained as suggested by the criteria above. We therefore do not condition on all possible household characteristics since they would likely explain away any differences between EITC recipients and non-recipients. Instead, we attempt to document the extent to which households that fit the EITC profile face borrowing constraints.

In Table 7, we report the means and standard deviations of these four measures for EITC recipients, non-recipients, and across household types. (Recall that EITC recipients in this context are defined as those who qualify for the EITC.) EITC recipient households report being denied a checking account because of bad credit more frequently than non-recipients (2.3 percent versus 0.5 percent for non-recipients). They also have lower credit card balances ($2,131 compared to $4,174); this could indicate that these households have lower credit limits, or are less willing to use acquire debt, or are less willing to use credit cards. EITC households are twice as likely to have late debt payments as non-recipients (11.2 percent compared to 5.4 percent), which would lead to having less access to credit. In addition, EITC households are three times more likely to not have a checking account (28 percent versus 7 percent).

When looking across households types, we can see that several interesting facts emerge. First, single households have lower credit card balances; they are generally more likely to have late payments; and they are less likely to have a checking account than married households (holding constant the number of children). However, the differences between single and married households are larger for non-recipients than for EITC recipients. For example, married households have much larger credit card balances than single households in the non-EITC sample, but the difference is smaller for married and single EITC recipients.

Second, married households with children that qualify for the EITC report very high late payment frequencies compared to their non-recipient counterparts. Approximately 13 percent of married households with one child have a late repayment, compared to just 5 percent of non-recipients. We do not observe significant differences between single-parent EITC recipients and non-recipients. Thus, EITC recipient households that are married with children will undoubtedly have worse credit statuses and lower borrowing limits than their non-recipient counterparts.

Third, for married households, credit does not seem to be more restricted for those with more children. However, single households seem to be more credit constrained as the number of children increases, and this is true for both

Table 7 Measures of Being Credit Constrained

7a: Measures of Being Credit Constrained: EITC Recipients vs. Non-Recipients

	Mean	**St. Dev.**
EITC Recipients:		
Bad Credit	2.3%	0.3%
Credit Card Balance (2007$)	$2,131	$140
Late Payment for 60+ Days	11.2%	0.6%
Has No Checking Account	27.9%	0.9%
Non-EITC Recipients:		
Bad Credit	0.5%	0.1%
Credit Card Balance (2007$)	$4,174	$91
Late Payment for 60+ Days	5.4%	0.2%
Has No Checking Account	7.0%	0.3%

7b: Measures of Being Credit Constrained by Household Type

	Married, No Kids	**Married, One Kid**	**Married, Two+ Kids**	**Single, No Kids**	**Single, One Kid**	**Single, Two+ Kids**
EITC Recipients:						
Bad Credit	0.0%	0.0%	2.2%	3.0%	3.3%	3.1%
Credit Card Balance (2007$)	$2,541	$2,966	$2,092	$2,419	$1,456	$1,933
Late Payment for 60+ Days	1.9%	13.6%	13.1%	7.0%	10.8%	13.5%
Has No Checking Account	28.7%	25.6%	25.7%	31.6%	27.3%	29.3%
Non-EITC Recipients:						
Bad Credit	0.2%	0.3%	0.0%	1.3%	0.0%	1.3%
Credit Card Balance (2007$)	$4,497	$4,946	$5,502	$2,693	$3,509	$1,401
Late Payment for 60+ Days	3.7%	3.9%	5.3%	6.5%	11.7%	8.5%
Has No Checking Account	5.0%	4.1%	2.1%	11.7%	13.2%	25.7%

Source: Authors' calculations using the 2007 SCF. Means are weighted.

EITC recipients and non-recipients. As documented above, the net worth of single households falls as the number of children increases (from Table 5a).

Our analysis suggests that EITC recipients use credit markets differently than non-recipients, possibly as a direct consequence of their income being currently and perhaps temporarily low, and this may have important implications on their ability to borrow. For example, EITC recipients are less likely to have a checking account and have lower credit card balances. They also more frequently have late debt repayments and are denied checking accounts than non-EITC recipients. Thus, it seems that at the time of receipt of the EITC, households are closer to limits on their ability to borrow than households that do not receive the EITC, and much of this is because of differences in income and household structure between the two groups.

9. CONCLUSION

In this article, we have studied several aspects of the Earned Income Tax Credit (EITC) that have been previously overlooked, including the income of EITC recipients at various ages, their wealth holdings, and the extent to which they are credit constrained. Naturally, we find that average annual earnings for those who receive the EITC are much lower than for non-EITC recipients at every age. In addition, younger households receive more EITC, and the amount of EITC received by these households suggests that the EITC increases lifetime earnings non-negligibly. The EITC in all likelihood provides a nontrivial mechanism for young, working households to smooth their consumption over their lifetimes.

The EITC acts as a negative income tax for recipient households. Specifically, we show that it has important implications on the marginal tax rate that low-income households face at various levels of earned income. Because of the phasing out of tax credits and income-support programs (such as TANF, food stamps, etc.), marginal income tax rates are much higher for low-income households than for middle- and high-income households in the United States. In particular, the marginal tax rate is negative for low levels of income, very high for those with moderate incomes that still qualify for the EITC, and then falls once households no longer qualify. We find that single-parent households that receive the EITC face some of the highest marginal income tax rates in the United States.

We then consider the theoretical and empirical effects of the EITC on the extensive and intensive margins of household labor supply. The EITC has undoubtedly increased labor force participation, but the effects on hours worked are ambiguous. This can be partly explained by the fact that low-income/low-wealth households that face borrowing constraints are insensitive to changes in the returns to working. Existing empirical work supports this conclusion.

Lastly, using data from the Survey of Consumer Finances, we estimate the wealth distribution of EITC households and measure the extent to which EITC households are credit constrained. Not surprisingly, we find that EITC-recipient households are very poor in terms of net worth: The average household has less than 20 percent of the average wealth of the average non-recipient household. In addition, EITC recipients are more likely to have bad credit and are more likely to have late debt payments than the average U.S. household, suggesting that they are more credit constrained.

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Economic Quarterly—Volume 96, Number 3—Third Quarter 2010—Pages 259–272

Instability and Indeterminacy in a Simple Search and Matching Model

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The search and matching model of Mortensen and Pissarides (1994) has become a popular and successful framework for analyzing labor market dynamics in dynamic stochastic general equilibrium (DSGE) models.[1] In this article, we point out a potentially problematic feature of this framework. We show that the solution to the dynamic model can be nonexistent or indeterminate. In particular, uniqueness problems arise when endogenous matching in response to labor market pressures is not elastic enough. In such a scenario, extraneous uncertainty, "sunspots," can lead to business cycle fluctuations even in the absence of any other disturbances. However, a solution does not exist when matching is too elastic. While these determinacy problems are plausible outcomes, we argue that they are not likely, as they are associated with regions of the parameter space that are at the extremes of typical calibrations.

Indeterminacy in search and matching models has previously been addressed by Giammarioli (2003). Her article differs from ours in that it introduces increasing returns in the matching function, which is a well-known mechanism to generate multiplicity in DSGE models (see Farmer and Guo [1994]). We show that indeterminacy in the search and matching model can arise even under constant returns. Our paper is similar to Burda and Weder

We are grateful to Andreas Hornstein, Anne Davlin, Marianna Kudlyak, and Alex Wolman for useful comments. The views expressed in this paper are those of the authors and should not necessarily be interpreted as those of the Federal Reserve Bank of Richmond or the Federal Reserve System. Krause is with the Economic Research Centre of the Deutsche Bundesbank in Frankfurt, Germany. Lubik is a senior economist with the Richmond Fed. E-mails: michael.u.krause@bundesbank.de; thomas.lubik@rich.frb.org.

[1] A nonexhaustive list of references includes Merz (1995); Andolfatto (1996); Cooley and Quadrini (1999); den Haan, Ramey, and Watson (2000); Krause and Lubik (2007); and Trigari (2009).

(2002) in this respect. Their indeterminacy results are driven, however, by the existence of labor market distortions, such as taxes, and the associated fiscal policy functions, and not by the features of the matching process per se. More recently, Hashimzade and Ortigueira (2005) analyzed the determinacy properties of a real search and matching model with capital. They show numerically how, for a given parameterization, the model admits sunspot equilibria. Zanetti (2006) incorporates the standard search and matching model into a New Keynesian DSGE model, where monetary policy is governed by an interest rate feedback rule. He shows that this expands the region of the parameter space where the Taylor principle, and thus equilibrium uniqueness, is violated. However, his paper focuses on the monetary policy rule as a source of indeterminacy. Labor market search and matching only provides a transmission mechanism, but is not analyzed as an independent factor of determinacy problems.

This article proceeds as follows. We present a canonical DSGE model with search and matching frictions in the next section. This is a bare-bones version of the model that does not rely on any increasing returns to scale in the functional forms. Our model specification has the advantage that the determinacy regions can be characterized largely analytically. Section 2 discusses issues related to the calibration of this model, while Section 3 derives its determinacy properties, both analytically and numerically. The final section briefly summarizes and concludes.

1. A CANONICAL DSGE MODEL OF LABOR MARKET SEARCH AND MATCHING

We develop a simple version of a discrete-time DSGE model with search and matching frictions in the labor market.[2] Key to the search and matching model is that new employment relationships are the result of time-consuming searches, both by firms and potential workers. In order to hire workers, firms first have to advertise open positions; they have to post vacancies, which is assumed to be costly. Existing matches between workers and firms are subject to job destruction, which leads to a flow of workers into the unemployment pool. The behavior of the aggregate economy is governed by the choices of a representative household, which engages in consumption smoothing. The household engages in perfect risk-sharing between its employed and unemployed members. The latter enjoy unemployment benefits while searching for a job. We employ some simplifying assumptions later on that lead to steady-state and dynamic equations that can be solved analytically. The properties of the full model are then analyzed numerically.

[2] The model is similar to Lubik (2009), to which we refer the reader for additional discussion and references.

Time is discrete. One period in the model is assumed to be a quarter. There is a continuum of identical firms that employ workers, who each inelastically supply one unit of labor.[3] Output, y, of a typical firm is linear in employment, n:

$$y_t = n_t. \tag{1}$$

The matching process is represented by a constant-returns matching function, $m(u_t, v_t) = m u_t^{\xi} v_t^{1-\xi}$, of unemployment, u, and vacancies, v, with parameters $m > 0$ and $0 < \xi < 1$. It captures the number of newly formed employment relationships that arise out of the contacts of unemployed workers and firms seeking to fill open positions. Unemployment is defined as

$$u_t = 1 - n_t, \tag{2}$$

which is the measure of all potential workers in the economy who are not employed at the beginning of the period and are thus available for job search activities.

Inflows to unemployment arise from exogenous job destruction at rate $0 < \rho < 1$. Employment therefore evolves according to

$$n_t = (1-\rho)[n_{t-1} + m(u_{t-1}, v_{t-1})]. \tag{3}$$

Note that newly matching workers who are separated from their job within the period reenter the matching pool immediately. We can define $q(\theta_t)$ as the probability of filling a vacancy, or the firm-matching rate, where $\theta_t = v_t/u_t$. We refer to θ as the degree of labor market tightness. In terms of the matching function, we can write this as $q(\theta_t) = m(u_t, v_t)/v_t = m\theta_t^{-\xi}$. Similarly, the probability of finding a job, the worker-matching rate, is $p(\theta_t) = m(u_t, v_t)/u_t = m\theta_t^{1-\xi}$. An individual firm is atomistic in the sense that it takes the aggregate matching rate, $q(\theta_t)$, as given. The employment constraint on the firm's decision problem is therefore

$$n_t = (1-\rho)[n_{t-1} + v_{t-1} q(\theta_{t-1})], \tag{4}$$

that is, it is linear in vacancy postings.

Firms maximize profits using the discount factor $\beta^t \frac{\lambda_t}{\lambda_0}$ (to be determined below):

$$\max_{\{v_t, n_t\}_{t=0}^{\infty}} \quad \sum_{t=0}^{\infty} \beta^t \frac{\lambda_t}{\lambda_0} [n_t - w_t n_t - \kappa v_t] + \\ + \sum_{t=0}^{\infty} \beta^t \frac{\lambda_t}{\lambda_0} \mu_t \left[(1-\rho)[n_{t-1} + v_{t-1} q(\theta_{t-1})] - n_t\right]. \tag{5}$$

[3] For expositional convenience, we present the problem of a representative firm only, and abstract from indexing the individual form and aggregation issues.

Wages paid to the workers are w, while $\kappa > 0$ is a firm's cost of opening a vacancy. μ is the Lagrange multiplier on the firm's employment constraint. It can be interpreted as the marginal value of a filled position. Firms decide how many vacancies to post (which can be turned into employment relationships) and how many workers to hire. The first-order conditions are

$$n_t \quad : \qquad \mu_t = 1 - w_t + \beta(1-\rho)\frac{\lambda_{t+1}}{\lambda_t}\mu_{t+1}, \tag{6}$$

$$v_t \quad : \qquad \kappa = \beta(1-\rho)\frac{\lambda_{t+1}}{\lambda_t}\mu_{t+1}q(\theta_t), \tag{7}$$

which imply a job-creation condition

$$\frac{\kappa}{q(\theta_t)} = (1-\rho)\beta\left(\frac{\lambda_{t+1}}{\lambda_t}\right)\left[1 - w_{t+1} + \frac{\kappa}{q(\theta_{t+1})}\right]. \tag{8}$$

This optimality condition trades off the expected hiring cost (which depends on the success probability $q(\theta_t)$) against the benefits of a productive match (which consists of the output accruing to the firms net of wage payments and the future savings on hiring costs when the current match is successful).

We assume that the economy is populated by a representative household. The household is composed of workers who are either unemployed or employed. If they are unemployed they are compelled to search for a job, but they can draw unemployment benefits, b. Employed members of the household receive pay, w, but share this with the unemployed. They do not suffer disutility from working and supply a fixed number of hours.[4] The household's only choice variable is consumption, so that its optimization problem is trivial:

$$\max_{\{C_t\}_{t=0}^{\infty}} \sum_{t=0}^{\infty} \beta^t \left[\frac{C_t^{1-\sigma} - 1}{1-\sigma}\right], \tag{9}$$

subject to

$$C_t = Y_t, \tag{10}$$

where C is consumption and Y is income earned from labor and residual profits from the firms; $0 < \beta < 1$ is the discount factor, and σ^{-1} is the intertemporal elasticity of substitution. From the household's (trivial) first-order condition we find that $\lambda_t = C_t^{-\sigma}$, where λ is the multiplier on the household's budget constraint. In equilibrium, total income accruing to the household equals net output in the economy, which is composed of production less real resources lost in the search process:

$$Y_t = y_t - \kappa v_t. \tag{11}$$

[4] We thus assume income pooling between employed and unemployed households and abstract from potential incentive problems concerning labor market search. This allows us to treat the labor market separate from the consumption choice. See Merz (1995) and Andolfatto (1996) for discussion of these issues.

Finally, we need to derive how wages are determined. We assume that wages are set according to the Nash bargaining solution.[5] Firms and workers maximize the bargaining function

$$\max_{w_t} (\mathcal{W}_t)^{\eta} (\mathcal{J}_t)^{1-\eta}, \tag{12}$$

with respect to the variable over which the two parties bargain, namely the wage, w_t. This results in the sharing rule:

$$\eta \mathcal{J}_t = (1-\eta) \mathcal{W}_t. \tag{13}$$

$\mathcal{W}_t$ denotes the match surplus accruing to the worker, while $\mathcal{J}_t$ is the firm's surplus, that is, the value of a filled job. The latter can be found from the firm's optimization problem. It is equal to the Lagrange multiplier on the employment constraint, μ_t, and is the shadow value of a filled position; to wit, $\mathcal{J}_t = \mu_t$. From the first-order condition with respect to employment we find that

$$\mathcal{J}_t = 1 - w_t + \beta(1-\rho)\frac{\lambda_{t+1}}{\lambda_t}\mathcal{J}_{t+1}. \tag{14}$$

The expression states that the value of a filled job is its marginal product, 1, net of wage payments, w_t; but it also has a continuation value $\mathcal{J}_{t+1}$, which is discounted at the time preference rate, β, and assuming that the filled job is still there next period. The latter is captured by the survival rate $(1-\rho)$.

We can derive the worker's surplus as follows. The worker receives payment in the form of the wage, w_t. But while he is working, he loses the value of being unemployed, b. The latter can be interpreted as the money value of enjoying leisure, engaging in household production, or simply unemployment benefits. Therefore, the current period net return is $w_t - b$. In the next period, the worker receives the continuation value $\mathcal{W}_{t+1}$, which is discounted at rate β. The worker has to take into account that he might not be employed next period, which is captured by the survival rate $(1-\rho)$, adjusted for the fact that a separated worker might not find a job again with probability $[1 - p(\theta_t)]$. Putting it all together, we have

$$\mathcal{W}_t = w_t - b + \beta(1-\rho)\left[1 - p(\theta_t)\right]\frac{\lambda_{t+1}}{\lambda_t}\mathcal{W}_{t+1}. \tag{15}$$

The two marginal values can now be substituted into the sharing rule and, after some algebra using the firm's first-order conditions, we can find the Nash-bargained wage:

$$w_t = \eta\left(1 + \kappa\theta_t\right) + (1-\eta)b. \tag{16}$$

[5] This is a standard assumption in the literature. Shimer (2005) provides further discussion.

We can now use this wage equation to derive the job-creation condition:

$$\frac{\kappa}{q(\theta_t)} = (1-\rho)\beta\frac{Y_t^\sigma}{Y_{t+1}^\sigma}\left[(1-\eta)(1-b) - \eta\kappa\theta_{t+1} + \frac{\kappa}{q(\theta_{t+1})}\right], \quad (17)$$

where we have used the first-order conditions of the household to eliminate the Lagrange multiplier, λ, from the discount factor. The dynamics of the model are given by the five equations in five unknowns: (2), (3), (11), (17), and the definition of labor market tightness, θ_t.

2. STEADY STATE AND CALIBRATION

We first compute the deterministic steady state of the model. We then linearize the dynamic system around the steady state and analyze the local determinacy properties of the economy. The equations describing the steady state are

$$\begin{aligned} u &= 1-n, && (18)\\ \theta &= \frac{v}{u}, && (19)\\ n &= \frac{1-\rho}{\rho} m v^{1-\xi} u^{\xi}, && (20)\\ Y &= n - \kappa v, && (21)\\ (1-\eta)(1-b) &= \frac{1-\beta(1-\rho)}{\beta(1-\rho)}\frac{\kappa}{m}\theta^{\xi} + c\eta\theta. && (22) \end{aligned}$$

(20) is the employment accumulation equation. It stipulates that inflows and outflows of the unemployment pool have to be equal. In a steady-state equilibrium, the number of separated workers, ρn, has to equal newly hired workers. Equation (22) is the job-creation condition, while the other equations are definitions.

There are five endogenous variables (u, n, v, θ, y) and seven structural parameters (ρ, m, ξ, κ, β, η, b). Because of the nonlinearity in the last equation, there is no analytical solution to this system. Given values for the parameters, however, we can compute a numerical solution. Using a nonlinear equation solver we determine θ from equation (22).[6] From equation (20) we can find $u = \left(1 + \frac{1-\rho}{\rho} m\theta^{1-\xi}\right)^{-1}$, and the solution for the other variables follows immediately.

We find it more convenient, however, to calibrate the model by fixing the steady-state unemployment rate, $u = \overline{u}$. This implies that one parameter has to be determined endogenously. Additionally, we can fix the endogenous

[6] Since the function in θ is monotonically increasing for nonnegative θ, there is a unique solution to this equation as long as $0 \leq b < 1$. This reflects the fact that the outside option of the worker, namely staying unemployed, cannot be larger than the worker's marginal product, i.e., the maximum rent that the worker can extract from the firm.

matching rate, $\overline{q} = m\theta^{-\xi}$, by using evidence on the rates at which firms fill vacancies. Hence, another parameter has to be determined endogenously. Using $n = 1 - \overline{u}$ in (20), we find that the match efficiency parameter is $m = \left(\frac{\rho}{1-\rho}\frac{1-\overline{u}}{\overline{u}}\right)^{\xi}\overline{q}^{1-\xi}$ and labor market tightness is $\theta = \left(\frac{m}{\overline{q}}\right)^{1/\xi}$. From (22) we can then also compute $\frac{1-b}{\kappa} = \frac{\eta}{1-\eta}\theta + \frac{1}{1-\eta}\frac{1-\beta(1-\rho)}{\beta(1-\rho)}\frac{\theta^{\xi}}{m}$. Note, however, that this condition does not pin down b and κ independently, nor does any other restriction in the model. Equation (21) helps only insofar as it restricts κ such that y remains positive. We chose to fix the vacancy cost parameter, κ, and let the benefit parameter, b, be determined endogenously.

For our calibration exercise we set the discount factor as $\beta = 0.99$. We chose a separation rate of $\rho = 0.1$. This is consistent with the evidence reported in Shimer (2005) and Lubik (2010), who use various econometric methods to estimate this parameter from U.S. labor market data. We agnostically set the bargaining parameter as $\eta = 0.5$ and follow most of the literature in this respect. Similarly, the match elasticity is $\xi = 0.5$, which is on the low end of estimates in the literature. Note that this benchmark calibration implements the Hosios condition, under which the market allocation in the model is socially efficient. The value for the match elasticity is at the low end of the plausible range as reported in the empirical study by Petrongolo and Pissarides (2001). We set the intertemporal substitution elasticity as $\sigma = 1$.

Finally, the two steady-state values are chosen as follows. We fix the unemployment rate, $\overline{u}$, at 12 percent. Our idea is to capture both measured unemployment in terms of recipients of unemployment benefits and potential job searchers that are only marginally attached to the labor force, but are open to job search. Since we do not model labor force participation decisions, this is a shortcut to capturing effective labor market search. This approach has been taken by Cooley and Quadrini (1999) and Trigari (2009). In choosing the steady-state job-matching rate, we follow den Haan, Ramey, and Watson (2000) who set $\overline{q} = 0.7$. In the numerical determinacy analysis below we conduct robustness checks for selected parameters and the calibrated steady-state values by varying them over their admissible range.

3. INDETERMINACY AND NONEXISTENCE

We now proceed by linearizing the dynamic equilibrium conditions around the steady state. It is a well-known feature of linear rational expectations models that they can have multiple equilibria, or that the solution may not even exist. We show that both scenarios are possible outcomes in the standard search and matching model, but they are associated with regions at the fringes of the parameter space. The linearized system is as follows (where $\widehat{x_t} = \log x_t - \log x$

denotes the percentage deviation of the variable x_t from its steady state x):

$$
\begin{aligned}
u\,\widehat{u}_t &= -n\,\widehat{n}_t, && (23)\\
\widehat{\theta}_t &= \widehat{v}_t - \widehat{u}_t, && (24)\\
\widehat{n}_{t+1} &= (1-\rho)\widehat{n}_t + \rho(1-\xi)\widehat{v}_t + \rho\xi\widehat{u}_t, && (25)\\
\widehat{Y}_t &= \frac{n}{y}\widehat{n}_t - \frac{\kappa v}{y}\widehat{v}_t, && (26)\\
\xi\widehat{\theta}_t - \sigma\widehat{Y}_t &= \left(\frac{\kappa\xi}{m}\theta^{\xi} - \eta\kappa\theta\right)X^{-1}\widehat{\theta}_{t+1} - \sigma\widehat{Y}_{t+1}, && (27)
\end{aligned}
$$

where $X = \frac{1}{\beta(1-\rho)}\frac{\kappa}{m}\theta^{\xi}$.

It is straightforward to substitute out $\widehat{u}_t$, $\widehat{v}_t$, and $\widehat{Y}_t$, so that we are left with

$$
\begin{bmatrix} \widehat{\theta}_{t+1} \\ \widehat{n}_{t+1} \end{bmatrix} = \begin{bmatrix} \frac{\xi+\sigma\frac{\kappa v}{y}}{\alpha_1} + \rho(1-\xi)\frac{\alpha_2}{\alpha_1} & -\frac{\alpha_2}{\alpha_1}\frac{\rho}{u} \\ \rho(1-\xi) & \frac{u-\rho}{u} \end{bmatrix} \begin{bmatrix} \widehat{\theta}_t \\ \widehat{n}_t \end{bmatrix}, \qquad (28)
$$

where $\alpha_1 = \beta(1-\rho)(\xi - \eta m\theta^{1-\xi}) + \sigma\frac{\kappa v}{y}$ and $\alpha_2 = \sigma\frac{n}{y}(1+\kappa\theta)$. This reduced form is expressed in terms of the state (or predetermined) variable, $\widehat{n}_t$, and the jump variable, $\widehat{\theta}_t$, which is a function of vacancy postings, $\widehat{v}_t$. The stability properties of the solution depend on the eigenvalues of the coefficient matrix. A unique solution requires that one root be inside the unit circle and the other root outside. Indeterminacy arises when both roots are inside the unit circle, while nonexistence occurs with both roots being explosive. In the former case, both equations are dynamically stable and an infinite number of paths (starting from arbitrary initial conditions) toward the unique steady state exist. In the latter case, both equations are explosive, which implies that, from any arbitrary initial condition, employment and vacancies would grow without bounds. This violates transversality or boundary conditions and can therefore not be an equilibrium.

The coefficient matrix is sufficiently complicated to prevent simple analytical derivations of the equilibrium regions. For illustrative purposes and for gaining intuition, we therefore make the simplifying assumption that the representative household is risk neutral, $\sigma = 0$. Later on, we discuss the general case using simulation results. Under risk neutrality, the coefficient matrix reduces to

$$
\begin{bmatrix} \frac{\xi}{\beta(1-\rho)(\xi-\eta p)} & 0 \\ \rho(1-\xi) & \frac{u-\rho}{u} \end{bmatrix}. \qquad (29)
$$

Since the coefficient matrix is triangular, the eigenvalues can be read off the principal diagonal. Recall that the worker matching rate is $p = m\theta^{1-\xi}$, which is equal to $\frac{\rho}{1-\rho}\frac{1-u}{u}$. Since we are treating the unemployment rate as a parameter to be calibrated, the determinacy conditions therefore only depend on structural parameters.

We establish the determinacy properties in the following proposition.

Proposition 1

1. *The model solution is indeterminate if and only if*
   - *(a)* $0 < \rho < 2u$,
   - *(b)* $0 < \xi < \frac{\beta(1-\rho)}{1+\beta(1-\rho)}\eta p$.
2. *The model solution is nonexistent if and only if*
   - *(a)* $\rho > 2u > 0$,
   - *(b)* $\frac{\beta(1-\rho)}{1+\beta(1-\rho)}\eta p < \xi < 1$.
3. *The model solution is unique if and only if either*
   - *(a)* $0 < \rho < 2u$,
   - *(b)* $\frac{\beta(1-\rho)}{1+\beta(1-\rho)}\eta p < \xi < 1$,

   or
   - *(c)* $\rho > 2u > 0$,
   - *(d)* $0 < \xi < \frac{\beta(1-\rho)}{1+\beta(1-\rho)}\eta p$.

Proof. Indeterminacy requires both roots inside the unit circle. Call $\lambda_2 = \frac{u-\rho}{u}$. It is straightforward to verify that $|\lambda_2| < 1$ over the permissible range iff $0 < \rho < 2u$. Call the other root $\lambda_1 = \frac{\xi}{\beta(1-\rho)(\xi-\eta p)}$. We have to distinguish two cases: if $\xi > \eta p$, no parameter combination can be found such that $|\lambda_1| < 1$. If $\xi < \eta p$, we can write $-\beta(1-\rho)(\xi-\eta p) > \xi > \beta(1-\rho)(\xi-\eta p)$. Simple algebra in combination with $\xi > 0$ then yields 1(b). Nonexistence requires that both roots be outside the unit circle. This is just the opposite scenario discussed before. Part 2 of the proposition follows immediately. Uniqueness requires one stable and one unstable eigenvalue. The parameter regions are consequently implied by those not considered in part 1 and 2. ■

The proposition shows that indeterminacy is a potential outcome in this model. It arises when the job destruction rate is less than twice the (calibrated) unemployment rate. For instance, at a separation rate of 10 percent, the unemployment rate would have to be less than 5 percent to definitely rule out indeterminacy on account of condition 1(a). This value is not implausible, given historical data for the United States where the average post-war unemployment rate is 4.8 percent. However, it has been argued (e.g., Trigari [2009]) that the proper corresponding concept for model unemployment includes not only the registered unemployed but also all workers potentially available for employment, such as discouraged workers or workers loosely attached to the labor force. Consequently, $\overline{u}$ should be assigned a much higher value (for instance, 26 percent as in Trigari [2009]), which raises the possibility of equilibrium indeterminacy.[7]

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[7] Calibrating $\overline{u}$ to a different value implies that benefits, b, and match efficiency, m, would have to change, too, since they are computed endogenously from the steady-state conditions. Higher

Condition 1(b) imposes an upper bound on the match elasticity, ξ. In the benchmark calibration, this upper bound is 0.147. Since ξ is typically calibrated to be above 0.5, this would rule out indeterminacy. However, this observation comes with the caveat that values for the match elasticity below 0.5 have some support in the literature. For instance, Cooley and Quadrini (1999) argue that a low elasticity in the range of $\xi = 0.1$ is necessary to match labor market cyclicality. Using likelihood-based econometric methods, Lubik (2010) finds that there is, in fact, substantial probability mass on low values of ξ. We also note that the upper bound is increasing in the Nash bargaining parameter. But even if $\eta \to 1$, indeterminacy would not occur for the typical parameter choices in the literature. Suppose, however, that the unemployment rate were set to $\overline{u} = 0.06$. In this case, the upper bound increases to 0.816, which would imply indeterminacy for typical search elasticity choices. Clearly, the interpretation of the pool of searchers in the matching model matters for determinacy questions.

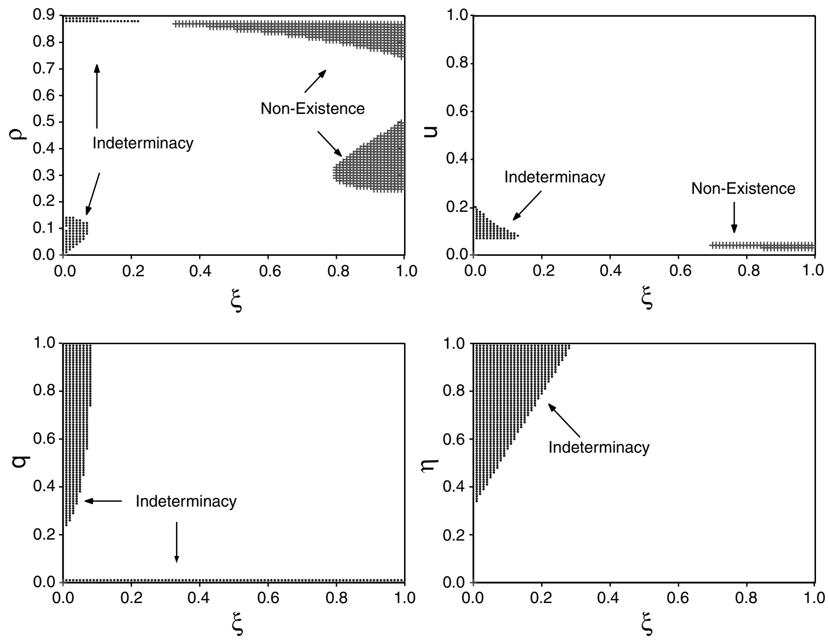
Intuitively, we can think about a sunspot equilibrium in the following way. Firms are willing to incur vacancy posting costs if they expect to recoup them through the proceeds from production net of wages and the savings on future hiring, as captured by the job-creation condition (17). The equilibrating mechanism is the behavior of the matching rate, $q(\theta)$. An increase in vacancy posting raises labor market tightness and lowers the probability that an individual firm is successful in finding an employee. This, in turn, raises effective hiring costs, $\kappa/q(\theta)$, which would have to be offset by higher expected returns. It is this externality, namely the fact that firms do not internalize the effect of their posting decisions on aggregate match probabilities, that is at the heart of the determinacy issue.[8]

Now suppose that a firm believes that future profits will be higher than is warranted by the fundamentals, such as the level of productivity. Beliefs of this kind can be triggered by sunspot shocks, as in the interpretation by Lubik and Schorfheide (2003). This belief would compel the firm to post more vacancies. If other firms were to do the same, aggregate tightness would increase and match probability would fall, raising effective hiring cost. What tends to rule out a sunspot equilibrium is that expected future benefits are not consistent with the higher posting costs. Consequently, rational firms do not act on sunspot beliefs. This argument breaks down in an environment where future benefits rise to accommodate higher current costs. The proposition

steady-state unemployment corresponds to a higher value of b and lower m. This can be interpreted as an implication of different labor market institutions.

[8] This has similar characteristics to the notion of an upward-sloping labor demand schedule in Farmer and Guo (1994). In their model, production exhibits constant returns to scale at the individual firm level, but increasing returns in the aggregate. An individual firm hiring more workers raises the marginal product of workers in the aggregate, thereby stimulating more labor demand. The job-creation condition can be thought of as a vacancy-demand curve.

Figure 1 Determinacy Regions



stipulates that indeterminacy arises when both the separation, ρ, and the match elasticity, ξ, are too small. When the former applies, the unemployment pool is small, while the latter makes new matches, and thereby future employment, highly elastic to vacancy postings. Consequently, the savings on future hiring costs react more than current effective costs, which helps validate sunspot beliefs.

A similar argument applies for the case of nonexistence of equilibrium. In general, nonexistence problems would arise for unemployment rates that are too low for given separation rates, in combination with excessively high match elasticities. In more technical terms, this combination makes the employment equation explosive. Any disturbance to a steady-state equilibrium would result in excessive job destruction (due to high separation rates) and matching that is inconsistent with the job-creation condition.

We now turn to the full model solution with risk-averse households ($\sigma >$ 0). We compute the determinacy regions numerically for combinations of the match elasticity, ξ, and various other structural parameters. The results are presented in Figure 1, where we have plotted determinacy regions for different subsets of the parameter space. The parameters are calibrated at the benchmark values discussed above. In each panel we vary two parameters over their admissible range while keeping the other parameters at their benchmark values.

As a general conclusion, determinacy problems tend to arise when the match elasticity, ξ, is either too small or too big. For small ξ, the equilibrium is indeterminate when the job destruction rate, ρ, or the unemployment rate, u, is too small. This is related to the analytical condition found in Proposition 1. Furthermore, firm-matching rates, q, above 0.2 and a Nash parameter that puts more weight on workers also lead to multiplicity. No equilibrium exists for large ξ and either a small unemployment rate or ρ above 0.2. We also analyze the sensitivity of the regions with respect to σ (not reported). As $\sigma \to 0$, the indeterminacy regions expand. In particular, any q implies multiple equilibria when $\xi < 0.2$. In the limit the boundaries between regions are given in the proposition. As the household becomes more risk averse, however, regions of indeterminacy disappear entirely.

An interesting special case to consider is a calibration with the Hosios condition, where $\eta = \xi$. This can be represented by a 45-degree line in the lower-right panel of Figure 1. In the absence of outside information on the value of the bargaining parameter, η, the Hosios calibration is often chosen in the literature. In this case, indeterminacy and nonexistence are ruled out and become highly unlikely for other parameter combinations. For instance, equilibrium nonexistence requires a separation rate of $\rho = 0.81$. Moreover, if $\eta = \xi$, we can rule out indeterminacy in the case of $\sigma = 0$ because condition 1(b) of the proposition never holds. The equilibrium could still be nonexistent, but this would require very high separation rates. In principle, these could obtain when the model period is much longer than a quarter since eventually all workers turn over within a long enough time horizon.[9]

Interpreting these results in light of standard calibrations used in the literature, we would argue that indeterminacy and nonexistence do not present serious problems for the search and matching framework. Hence, it is unlikely that sunspot equilibria would be helpful in explaining labor market dynamics (as claimed in Hashimzade and Ortigueira [2005]). This is not to say that labor search and matching frameworks cannot support indeterminate equilibria. Mildly increasing returns to scale in the matching function (Giammarioli 2003) lead to widely expanded indeterminacy regions, while a New Keynesian model with search and matching frictions in the labor market has broader indeterminacy properties than the standard New Keynesian model (Zanetti 2006).

[9] Incidentally, the continuous-time version of this simple search and matching model always has a unique solution (see Shimer [2005]), as the separation-relevant time horizon is infinitesimally small. We are grateful to Andreas Hornstein for pointing this out.

4. CONCLUSION

We show in this article that for most plausible parameterizations the simple search and matching model does not suffer from determinacy problems. Specifically, we argue that it is unlikely that the model has multiple equilibria so that extraneous uncertainty, i.e., animal spirits, can cause business cycles. Parameterizations that lead to indeterminacy can be found, but they lie at the boundaries of the region that the empirical literature would consider plausible. We identify the match elasticity and the separate rate as crucial parameters in that respect.

These properties are obviously model specific, but our conclusions are likely robust to modifications such as endogenous job destruction. While the boundaries of the determinacy regions are likely to shift, the dynamic mechanism stays unaffected. The main caveat to our study is that our analysis applies to a local equilibrium in the neighborhood of the steady state. However, the underlying model is nonlinear and local results may therefore not adequately describe the global equilibrium properties. Naturally, this is a topic for further investigation. Moreover, researchers may actually be interested in the business cycle implications of indeterminacy that do not depend on policy rules or externalities. It appears plausible that actual labor market decisions are characterized to some extent by animal spirits. Further research should shed some light on this issue.

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Economic Quarterly—Volume 96, Number 3—Third Quarter 2010—Pages 273–290

How Large Has the Federal Financial Safety Net Become?

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In 2002, Walter and Weinberg examined the federal financial safety net as it stood at the end of 1999 (Walter and Weinberg 2002). At the time, the authors estimated that approximately 45 percent of all financial firm liabilities were protected by the safety net. As one would expect in this article, the current estimate indicates that the size of the net has grown, as the financial market turmoil that began in 2007 led federal government agencies to expand the range of institutions and the types of liabilities protected by the safety net.

1. THE SAFETY NET: ITS DEFINITION, COSTS, AND BENEFITS

Walter and Weinberg defined the federal financial safety net as consisting of all explicit or implicit government guarantees of private financial liabilities. Private financial liabilities are those owed by one private market participant to another. As used by Walter and Weinberg, the phrase *government guarantee* means a federal government commitment to protect lenders from losses due to a borrower's default (Walter and Weinberg 2002).[1] Following this definition, we include in our estimate of the safety net, insured bank and thrift deposits, certain other banking company liabilities, some government-sponsored enterprise (GSE) liabilities, selected private employer pension liabilities, as well as

The authors would like to thank Jason Annis, Marc Chumney, Tim Pudner, and Deanna West for providing data and valuable advice, as well as Huberto Ennis, Robert Hetzel, Sabrina Pellerin, and John Weinberg for their insightful comments on an earlier draft. The views expressed in this article are those of the authors and do not necessarily reflect those of the Federal Reserve Bank of Richmond or the Federal Reserve System. E-mail: john.walter@rich.frb.org.

[1] In addition to estimating the proportion of financial firm liabilities backed by the federal government, Walter and Weinberg also estimated the proportion of nonfinancial firm and household liabilities with such backing.

a subset of the liabilities of other financial firms. The details of why we chose to include these liabilities are provided below.

Effect of a Safety Net on Economic Efficiency

Government actions in the form of subsidies, taxes, or regulations change market outcomes, and in competitive markets such changes distort allocations and can reduce economic efficiency. Does the financial safety net cause distortions? As discussed in Walter and Weinberg, in principle, the government could design guarantees that mimic market outcomes. Typically, however, government intervention arises from a desire to alter market outcomes. In the case of guarantees, this means either expanding coverage or underpricing relative to private market guarantees. Underpricing means that the guarantor collects fees that are less than the expected value of its obligations. This underpricing subsidizes risk taking.

Underpriced guarantees tend to shift resources away from activities that are not covered toward those that are. In that way, a government guarantee is similar to a direct subsidy paid to those engaged in a particular activity. A guarantee is different, however, in the way it affects attitudes toward risk. By assigning to the government part of the risk in the activities being financed, the safety net reduces market participants' willingness to control risk. Overprovision of guarantees, while not necessarily drawing resources into an activity, does shift risk preferences in a way similar to underpricing. In short, guarantees lead to expanded risk taking.

Our calculation of the size of the safety net does not represent a measure of the size of the distortions to the allocation of resources and risk taking. Such a measure would require knowledge of the extent of underpricing or overprovision of government guarantees. Those would be difficult to measure, especially the latter, since government provision often preempts private market activity. We nevertheless believe that the extent of distortions is directly related to the size of the safety net. Other things being equal, the greater the share of private liabilities protected by the government safety net, the more likely it is that government guarantees are extending beyond the level of protection that would be provided in a private market.

Why Have a Safety Net?

If the safety net is distortionary, why have one? Proponents of the financial safety net, especially as it applies to banks, often argue that private risk-sharing arrangements tend to disregard the *systemic* consequences of large losses borne by an individual or a small group of institutions. The idea here is that such losses might spill over and generate further losses caused, for example, by a contagious loss of investor confidence. Under such a view, government protection for certain investors could prevent widespread financial panic or distress. While the potential systemic consequences of a large financial failure are difficult to assess, when faced with the possibility of widespread failures of financial firms, policymakers are likely to conclude that preventing such failures by protecting creditors of financial firms (providing safety net protection) is prudent.

Similarly, some observers maintain that the safety net protections can lower the costs of, and therefore encourage, certain highly beneficial financial arrangements. For example, Diamond and Dybvig (1983) argue that banks' performance of the *maturity transformation function* is highly beneficial to the economy but is more costly without government-provided deposit insurance. Banks perform maturity transformation by gathering money from numerous short-term depositors (those bank customers whose deposits *mature* soon after deposited—especially checking deposits, which are available, meaning that they mature, immediately after being deposited) to fund long-term loans to businesses and individuals. Without deposit insurance, which only the government has sufficient resources to provide, bank runs are likely to occur. A bank run happens when many depositors attempt to withdraw their funds simultaneously. Since banks make long-term loans, they cannot recover sufficient money from borrowers to meet a run and, therefore, fail. To protect themselves from runs, banks can undertake costly private measures, but Diamond and Dybvig argue that government deposit insurance is likely to be less expensive and therefore preferable to such measures.

2. LEGISLATIVE AND REGULATORY CHANGES THAT EXPANDED THE SAFETY NET

As shown in Table 1, we estimated the proportion of financial firm liabilities protected as of the end of 2009. By the end of 2009, a number of government programs had been established to address turmoil in financial markets. Employing methods similar to those used by Walter and Weinberg when they measured the size of the safety net for the end of 1999, we find that as of the end of 2009 about 59 percent of financial firm liabilities were protected by the federal safety net.

One of the most important reasons for the increase from 1999 to 2009 is the enlarged portion of banking firm liabilities that market participants are likely to consider protected: banking and savings firm liabilities with an implicit backing. In 1999, implicitly guaranteed liabilities of banks and savings institutions amounted to about 13 percent of all of these firms' liabilities (15.9 percent for commercial banks and 4.2 percent for savings institutions), or \$820

Table 1 Estimated Federal Financial Safety Net

Financial Firms	Explicitly Guaranteed Liabilities	Implicitly Guaranteed Liabilities	Explicitly and Implicitly Guaranteed Liabilities	Total Liabilities
Banking and Savings Firms				
(Includes BHCs)	6,536	7,276	13,812	16,249
	40.2%	44.8%	85.0%	
Credit Unions	725		725	817
	88.7%		88.7%	
Government-Sponsored Enterprises				
Fannie Mae		3,345	3,345	3,345
Freddie Mac		2,333	2,333	2,333
Farm Credit System		188	188	188
Federal Home Loan Banks		973	973	973
Total		6,838	6,838	6,838
		100%	100%	
Private Employer Pension Funds	2,799		2,799	3,273
	85.5%		85.5%	
Other Financial Firms		748	748	15,158
		4.9%	4.9%	
Total for Financial Firms	10,059	14,862	24,921	42,335
	23.8%	35.1%	58.9%	

Notes: Data from December 2009, in billions of dollars. Figures may not sum exactly due to rounding. The figures in the column “Explicitly and Implicitly Guaranteed Liabilities” are the sum of the numbers in the first two columns, “Explicitly Guaranteed Liabilities” and “Implicitly Guaranteed Liabilities.” See Appendix for table legend.

billion; in 2009, about 45 percent of banking and savings firm liabilities were implicitly guaranteed, by our estimate, amounting to $7.3 billion.[2]

How did Walter and Weinberg determine which institutions to include as having an implicit guarantee and which liabilities issued by these institutions might be covered? As the authors noted, the critical question is whether market participants believe that a given institution will be protected, even though official policy may not state explicitly that all of these liabilities are protected. As of 1999, Walter and Weinberg argued that market participants were likely to assume that certain holders of liabilities in the largest 21 banking companies and the two largest thrift companies would be protected in the event that these firms became troubled. These 21 banking companies and two thrifts all had assets (in 1999 dollars) of more than $50 billion, which was greater than the smallest of the 11 institutions identified by the Comptroller of the Currency in 1984 as potentially too big to fail (Walter and Weinberg 2002, p. 381). The liabilities that Walter and Weinberg assumed the market would be highly likely to view as protected were deposits of more than $100,000 (deposits of less than $100,000 are included in the "Explicitly Guaranteed Liabilities" column in the tables), federal funds loans made to the 21 banks and two thrifts, and repo transactions with these banks and thrifts. Though we intend to use a similar methodology for estimating the size of implicit guarantees for banking companies in 2009, events during the recent financial crisis required some adjustments.

Support for Stress-Tested Financial Companies

Given that the government had responded aggressively to problems in financial firms during the financial turmoil of 2008–2009, our challenge is to decide which institutions have implicit guarantees. Here we maintain that market participants were very likely to assume that the liabilities of the financial firms that were stress tested early in 2009 (participants in the Supervisory Capital Assessment Program—SCAP) had a strong likelihood of receiving federal backing if they suffered financial distress. Indeed, the announcement of the stress tests in February 2009 came with a promise of government-provided capital for stress-tested institutions that were shown to be in need of additional capital:

> Under [the Treasury's Capital Assistance Program] CAP, federal banking supervisors will conduct forward-looking assessments [SCAP stress tests] to evaluate the capital needs of the major U.S. banking institutions under a more challenging economic environment. Should that assessment indicate that an additional capital buffer is warranted, banks will have

[2] An explanation of the factors underlying the large increase is provided below.

> an opportunity to turn first to private sources of capital. In light of the current challenging market environment, the Treasury is making government capital available immediately through the CAP to eligible banking institutions to provide this buffer. (FinancialStability.gov 2009)

Additionally, a number of these firms did, in fact, receive government aid in the form of capital injections in 2008 and early 2009 through the Treasury's Capital Purchase Program or in response to the stress tests (FinancialStability.gov 2010, pp. 21, 27, 67–80). This aid, both the aid promised under the CAP and aid received through the Capital Purchase Program, reduced the likelihood that *all* liabilityholders of the protected firms would suffer losses, so here we include *all* liabilities of the stress-tested banking institutions in our safety net calculation.

While some observers in 2009 may have viewed the likely passage of financial reform legislation as diminishing federal backing, we nevertheless count the liabilities of the stress-tested firms. Legislation that was intended to limit the chance that financial institutions would receive federal aid was being considered in the U.S. Congress during 2009. If market participants were convinced that such legislation would forestall any opportunity for the creditors of the largest financial institutions to be protected by the federal government, then our calculation might appropriately exclude the liabilities of stress-tested banking institutions. In fact, most of the legislative proposals included language that called for the closure of troubled financial firms with losses to equityholders and at least some creditors (though at least one leading proposal contained protections for creditors of financial firms if the failure of such a firm might create a systemic risk).[3] Nevertheless, legislative proposals contained provisions meant to establish a mechanism that could clearly identify "systemically important" financial firms. Such mechanisms seem likely to encourage market participant expectations of federal aid to the creditors of the largest (i.e., systemically important) firms. Given the ambiguous effect of the reform proposals on the probability of federal aid to the largest banking firms, and the clear protections provided for troubled firms and for their creditors during the financial turmoil, we retain their liabilities in our estimate of liabilities protected by the safety net, in keeping with Walter and Weinberg (2002). (In a later section we remove the liabilities of stress-tested institutions and re-estimate the size of the safety net—see Table 2.)

As indicated earlier, the total liabilities of the 19 stress-tested bank holding companies, less their liabilities that were explicitly covered by deposit insurance, summed to $7.3 trillion ("Implicitly Guaranteed Liabilities" column in

[3] See H.R. 4173 as of December 2, 2009, p. 370, available at: http://www.house.gov/apps/list/press/financialsvcs_dem/presscfpa_121109.shtml.

the tables). This sum equals about 45 percent of all banking and savings firm liabilities.

Increased Ceiling on Insured Deposits

Several Federal Deposit Insurance Corporation (FDIC) programs expanded the explicit portion of the safety net for banks and thrifts ("Explicitly Guaranteed Liabilities" column in the tables) beyond the long-standing $100,000 coverage for deposits (which are also included in the "Explicitly Guaranteed Liabilities" column in the tables).[4] For example, in October 2008 the Emergency Economic Stabilization Act of 2008 temporarily increased FDIC deposit insurance coverage from $100,000 to $250,000, until December 31, 2009. In May 2009, the $250,000 cap was extended to December 31, 2010, by the Helping Families Save Their Homes Act. In July 2010, legislation made permanent the $250,000 coverage limit (Federal Deposit Insurance Corporation 2010a).

Transaction Account Guarantee Program

Further, in October 2008 the FDIC implemented a program to insure uninsured deposits (those deposits in accounts containing more than $250,000) in noninterest-bearing transactions accounts for those insured banks and thrifts wishing to participate. The program is temporary. At first it covered such transactions accounts until December 31, 2009. Later the FDIC extended the program's coverage until June 30, 2010, and then extended it again until December 31, 2010, with a pre-announced option to extend it an additional 12 months (Federal Deposit Insurance Corporation 2010a).[5] This program, the Transaction Account Guarantee Program (TAGP), added $834 billion to our "Explicitly Guaranteed Liabilities" column in the tables for banking and savings firms (Federal Deposit Insurance Corporation 2009c).

Debt Guarantee Program

Last, in October 2008 the FDIC offered, to banking and savings institutions wishing to participate, the option to receive FDIC insurance coverage for senior unsecured debt issued by such institutions. This Debt Guarantee Program

[4] Since April 2006, deposits in certain retirement accounts at banks and thrifts have been protected by the FDIC up to $250,000 (Federal Deposit Insurance Corporation 2006). Deposits in such accounts, up to the $250,000 ceiling, are included in the "Explicitly Guaranteed Liabilities" column of our tables.

[5] The Dodd-Frank Wall Street Reform and Consumer Protection Act extended coverage for noninterest-bearing transaction accounts through December 31, 2012 (Federal Deposit Insurance Corporation 2010c).

(DGP) at first covered debt issued by June 30, 2009, and maturing by June 30, 2010. The DGP was later extended to cover debt issued by October 31, 2009, and maturing by December 31, 2012. As of December 31, 2009, the program was insuring $309 billion in debt (Federal Deposit Insurance Corporation 2009b).

3. OTHER COMPONENTS OF THE SAFETY NET

As in 1999, we include for 2009 the liabilities of government-sponsored enterprises (direct GSE liabilities plus the dollar amount of mortgage-backed security guarantees) in the "Implicitly Guaranteed Liabilities" column in the tables. Earlier we noted that government guarantees can often modify market prices. Though our article has made no attempt to measure the size of guarantees' effect on market prices, in the case of the GSEs' implicit guarantee, the size of the effect on market prices has been estimated by Passmore (2005) and others.[6] Passmore (2005) estimates that the average homeowner saved between 3 and 11 basis points on his or her mortgage because of the implicit guarantee. The subsidy lowers the GSEs' borrowing costs, and some of this saved borrowing cost is passed on to homeowners by the GSE in the form of lowered mortgage interest rates. Passmore calculates that about half of the guarantee's benefit flows to the shareholders of the GSEs. While the Treasury made clear its support for Fannie Mae and Freddie Mac once these two financial firms were placed in conservatorship in September 2008, the support was not as strongly stated as that given to insured deposits, so we leave these liabilities in the implicit column in the tables.[7]

We estimate the amount of private pensions explicitly guaranteed in 2009 by the Pension Benefit Guarantee Corporation (PBGC) based on the latest private pension data available, which are data for 2007 (Pension Benefit Guarantee Corporation 2010, pp. 83, 105). Our admittedly rough 2009 figure is derived by simply adjusting the 2007 figure by twice the average annual growth rate of private pension liabilities for the previous 10 years (1997–2007).

We also count all of the liabilities of American International Group (AIG) as implicitly guaranteed in the "Other Financial Firms" row in the tables.[8]

[6] Beyond Passmore, the Congressional Budget Office (2001) also developed estimates of the GSEs' guarantee on mortgage interest rates.

[7] We treat Fannie Mae and Freddie Mac as private entities and therefore include their liabilities in our table, consistent with the way Walter and Weinberg treated these entities, even though the status of Fannie Mae and Freddie Mac as privately owned firms is more ambiguous now than in 1999.

[8] The insured deposit liabilities of AIG's savings bank are not included in the "Other Financial Firms" row since these liabilities were included in the "Banking and Savings Firms" row. While AIG owns a savings bank, it is not classified as a bank holding company (and does not file a bank holding company report [Y9C] with federal regulators), so we do not include it in the Banking and Savings Firms row.

We count their liabilities as such because of the aid provided them by the Federal Reserve and the U.S. Treasury following AIG's financial problems in September 2008. Because there were no clear signals about whether aid might be forthcoming for other large, nonbank financial firms (beyond the stress test firms), we did not include the liabilities of any firms other than AIG in the "Other Financial Firms" row in tables.

4. ALTERNATIVE ESTIMATES OF THE SIZE OF THE SAFETY NET

As has been noted, Table 1 is based on several assumptions similar to those made by Walter and Weinberg in 2002. For example, we assumed that all liabilities of stress-tested bank holding companies would be protected, not just the liabilities representing FDIC-insured bank deposits. What would be the size of the safety net if these assumptions were changed?

Contrary to our assumption about the likely protection of liabilityholders of stress-tested companies, one can imagine circumstances under which such liabilityholders might be left unprotected. If one of these companies were to fail at a time when financial markets were broadly healthy, policymakers could more easily allow the company to be handled as a bankruptcy so that no government funds are employed to protect liabilityholders (of course, the holders of FDIC-insured deposits would still be covered given that such deposits are protected regardless of the circumstances surrounding the failure). In times of general financial market strength, the failure of a large holding company could perhaps be absorbed without worries of a cascade of additional failures. And at such times, if the firm were handled through the Dodd-Frank Act's orderly liquidation process, it is possible that neither the government nor other financial firms would provide funds to protect liabilityholders.[9]

While investors might expect large financial firm failures to typically occur in times of widespread financial weakness, and therefore anticipate that their investments would be protected, some large firms have failed in times of financial market health. One such example was London-based Barings Bank, which failed when financial markets were broadly strong in 1995. Its failure was because of the huge trading losses generated by one unchecked Barings trader who took large, unauthorized futures positions. Given that there are circumstances under which the holders of stress-tested company liabilities might be left unprotected, dropping the assumption of their coverage and recalculating our estimate of implicitly guaranteed liabilities seems worthwhile.

[9] The Orderly Liquidation Authority section of the Dodd-Frank Wall Street Reform and Consumer Protection Act of 2010 contains provisions that allow funds gathered from assessments on the largest financial firms to be used to protect liabilityholders.

Table 2 Estimated Federal Financial Safety Net, Narrowly Defined

Financial Firms	**Explicitly Guaranteed Liabilities**	**Implicitly Guaranteed Liabilities**	**Explicitly and Implicitly Guaranteed Liabilities**	**Total Liabilities**
Banking and Savings Firms				
(Includes BHCs)	5,392		5,392	16,249
	33.2%		33.2%	
Credit Unions	725		725	817
	88.7%		88.7%	
Government-Sponsored Enterprises				
Fannie Mae		3,345	3,345	3,345
Freddie Mac		2,333	2,333	2,333
Farm Credit System		188	188	188
Federal Home Loan Banks		973	973	973
Total		6,838	6,838	6,838
		100%	100%	
Private Employer Pension Funds	2,799		2,799	3,273
	85.5%		85.5%	
Other Financial Firms				15,158
Total for Financial Firms	8,915	6,838	15,753	42,335
	21.1%	16.2%	37.2%	

Notes: Data from December 2009, in billions of dollars. Figures may not sum exactly due to rounding. The figures in the column "Explicitly and Implicitly Guaranteed Liabilities" are the sum of the numbers in the first two columns, "Explicitly Guaranteed Liabilities" and "Implicitly Guaranteed Liabilities." See Appendix for table legend.

Large financial firms that are not bank holding companies might receive no protection in such instances, so we also drop liabilities of AIG from those liabilities with implicit backing.

Also, we included in our explicitly insured deposits category those deposits covered by the FDIC's temporary guarantee programs, since these programs were in place in 2009. But under the debt guarantee program no new debt issues were covered after October 31, 2009 (Federal Deposit Insurance Corporation 2010b). The TAGP was set to expire as of the end of 2010, though the Dodd-Frank Act extended it to December 31, 2012. In the case of future financial firm failures, such programs may not be in place, and might not be reinstated. Therefore, re-estimating our measure of the size of the safety net without considering these deposits as protected also seems worthwhile.

Table 2 contains our estimate of the size of the safety net without including the liabilities of the stress-tested bank holding companies, AIG, and the FDIC temporary insurance program deposits. These changes mean that, compared to Table 1, the proportion of liabilities receiving explicit and implicit guarantees falls to 37.2 percent.

Additionally, while we assume that the liabilityholders of the housing and farm credit GSEs will be protected from loss, as were such holders of Fannie Mae and Freddie Mac debt during the 2007–2009 financial crisis, under some circumstances such holders might be left unprotected. As in the case of the stress-tested companies, if a GSE were to fail during a period in which financial markets were healthy, policymakers might leave debtholders unprotected. Therefore, it is possible that one might want to exclude the liabilities of the GSEs from the calculation of the safety net. If the $6.8 trillion in liabilities of the GSEs were removed (which are the only implicitly guaranteed liabilities in Table 2), then our measure of the safety net would shrink to 21 percent of total liabilities in Table 2, the amount of explicit liabilities shown in Table 2.

Some readers might contend that one category of liabilities, which we have excluded from our safety net estimate, could legitimately be added: money market mutual fund liabilities. In the creation of our tables, and in Walter and Weinberg (2002), mutual fund liabilities are excluded because the principal value of mutual fund investments, including money market mutual fund investments, can decline, without the mutual fund defaulting, if the entity in which the funds are invested defaults. As a result, these investments are akin to equity and unlike private liabilities—the focus of our estimates—which typically must pay back full principal (or else be in default). For example, an investor in a money market mutual fund, which in turn invested in financial firm commercial paper, could lose principal if the commercial paper was not repaid, but the mutual fund can continue to operate (i.e., not default).[10] This

[10] Money market mutual funds are loath to pay back less than full principal ("break the buck" in mutual fund parlance), and few have done so over time. Instead, the money market

view of money market mutual fund investments as equity must be tempered, however, by events in 2008. Specifically, the Treasury stepped in and protected investors in mutual funds from losses, thereby treating investments in the funds like other guaranteed *liabilities*, in which losses are prevented by government assistance or guarantees. As a result, one might argue that our estimates of the fraction of total liabilities carrying a government guarantee—both the numerator and denominator—should include money market mutual funds. If one adds the amount of such fund balances outstanding at the end of 2009 ($3.3 trillion [Investment Company Institute 2010]) to our estimates in the column "Explicitly and Implicitly Guaranteed Liabilities" in Table 1, the proportion would increase to 62 percent. The Table 2 figure would increase to 42 percent.

5. CONCLUSION

Recent government actions by legislators and financial regulators expanded the federal financial safety net. Such actions include augmentation of deposit insurance, debt guarantees for banking companies, aid to stress-tested financial firms, and, perhaps, various regulatory reform legislative proposals. As discussed in Walter and Weinberg (2002), this expansion has likely encouraged a view that liabilityholders will be protected by the federal government in times of financial difficulty in the future. As a result of this expectation of government protection, liabilityholders will exercise less oversight over financial firm risk taking then they would without this expectation, financial firms will undertake more risk, and financial market decisions will be distorted and inefficient.

mutual fund's parent typically injects funds to allow the fund to pay back full principal. This behavior by mutual fund parent companies indicates that parent companies and investors may well view money market mutual fund investments more as liabilities than equity, regardless of the fact that money market mutual funds can break the buck without defaulting.

APPENDIX A: LEGEND TO TABLE 1

- Banking and Savings Firms[11]
  - Explicitly Guaranteed Liabilities
    - * FDIC-insured deposits of all commercial banks and savings institutions including transaction accounts covered by the FDIC's TAGP, plus debt guaranteed by the FDIC's DGP
  - Implicitly Guaranteed Liabilities
    - * Total liabilities of the 19 stress-tested institutions, less FDIC-insured deposits and accounts covered by TAGP and debt covered by DGP for the 19 stress-tested institutions
- Credit Unions
  - Explicitly Guaranteed Liabilities
    - * National Credit Union Administration-insured shares and deposits
- Government Sponsored Enterprises
  - Implicitly Guaranteed Liabilities of:
    - * Fannie Mae
      - · Total liabilities
      - · Fannie Mae mortgage-backed securities held by third parties
      - · Other guarantees
    - * Freddie Mac
      - · Total liabilities
      - · Freddie Mac participation certificates and structured securities held by third parties
    - * Farm Credit System
      - · Total liabilities
      - · Farmer Mac guarantees
    - * Federal Home Loan Banks
      - · Total liabilities

[11] See Section 4 for a description of the differences between Table 1 and Table 2 estimates.

- Private Employer Pension Funds
  - Explicitly Guaranteed Liabilities
    - * Pension liabilities backed by the PBGC
- Other Financial Firms
  - Explicitly Guaranteed Liabilities
    - * Total liabilities of AIG, less FDIC-insured deposits of AIG Federal Savings Bank

APPENDIX B: DATA APPENDIX TO TABLE 1

Banking and Savings Firms—Explicitly Guaranteed Liabilities:

"Estimated FDIC-insured deposits" of commercial banks, savings institutions, and U.S. branches of foreign banks (Federal Deposit Insurance Corporation 2009a), plus "Amount Guaranteed" in the Transaction Account Guarantee Program (Federal Deposit Insurance Corporation 2009c), plus "Debt Outstanding" in the Debt Guarantee Program (Federal Deposit Insurance Corporation 2009b).

Banking and Savings Firms—Implicitly Guaranteed Liabilities:

Total liabilities of the 19 stress-tested institutions found in the Y9C (quarterly bank holding company financial reports), less 1) the explicitly guaranteed deposits of the banks and savings institutions owned by these 19 firms, and 2) the FDIC-insured debt (insured under the DGP) of each of these institutions. The estimated FDIC-insured deposits and the guaranteed amount in noninterest-bearing transaction accounts for each bank can be found on the FDIC's website in the "Institution Directory" (www2.fdic.gov/idasp). The amount of DGP debt of each firm can be found on the firms' 10Ks.

Banking and Savings Firms—Total Liabilities:

Total liabilities from the following sources: For large (consolidated assets of over $500 million) bank holding companies, Consolidated Financial Statements for Bank Holding Companies (FR Y9C); for small (consolidated assets less than $500 million) bank holding companies, Parent Company Only Financial Statements for Small Bank Holding Companies (FR Y9SP)—from which consolidated total liabilities can be derived; for banks not owned by a bank holding company, Consolidated Reports of Condition and Income for a Bank (FFIEC 031 and FFIEC 041); and for all thrift liabilities, Thrift Financial Reports.

Credit Unions—Explicitly Guaranteed Liabilities:

Total insured shares at the $250,000 limit (National Credit Union Administration 2009).

Credit Unions—Total Liabilities:

Board of Governors (2010), Table L.115—Credit Unions, "Total liabilities."

Government-Sponsored Enterprises:

Fannie Mae:

Total liabilities, plus Fannie Mae MBS held by third parties, plus other guarantees found in the Fannie Mae 10K, "Item 6. Selected Financial Data" (p. 70).

Freddie Mac:

10K report of Freddie Mac, "Total liabilities" ("Consolidated Balance Sheets," p. 209), plus "Total PCs and Structured Securities issued" ("Item 6. Selected Financial Data," p. 57), less "Total Freddie Mac PCs and Structured Securities held" in Freddie Mac portfolio (Table 28, p. 104).

Farm Credit System:

Farm Credit System (2010), "Total liabilities" ("Combined Statement of Condition Data," p. 3), plus "Farmer Mac guarantees" (p. 12).

Federal Home Loan Banks:

Federal Home Loan Banks (2010), "Total liabilities" ("Combined Statement of Condition," p. 194).

Private Employer Pension Funds—Explicitly Guaranteed Liabilities:

Liabilities of all pension funds insured by the PBGC (which insures only defined benefit plans) were $2,559 billion in 2007, the latest date for which data are reported (Pension Benefit Guarantee Corporation

2010, pp. 83, 105). This figure is inflated by twice (because 2007–2009 involves two years of growth) the average annual growth rate of PBGC-insured pension liabilities from 1997–2007 to obtain our estimate of all liabilities in pension funds insured by the PBGC as of December 31, 2009 ($2,946 billion). Since PBGC covers pensions only up to a specified maximum payment per year, a portion of beneficiaries' pensions in guaranteed plans—those with pensions paying above this maximum—are not insured. According to the PBGC, this portion is estimated to be 4–5 percent (Pension Benefit Guarantee Corporation 2007, p. 24; Pension Benefit Guarantee Corporation 1997, footnote to Table B-5). To arrive at the guaranteed portion of PBGC guaranteed pension fund liabilities, we multiplied total 2009 fund liabilities ($2,946 billion) by 0.95 to yield $2,799 billion.

Private Employer Pension Funds—Total Liabilities:

There appears to be no data on the total liabilities of all private employer-defined benefit pension funds. Therefore, we estimate our total liability figure based on PBGC data. To derive our figure, we begin with our previously determined estimate of all private pension fund liabilities that are included in PBGC ($2,946) and then divide it by 0.9 to arrive at our total liability figure of $3,273 billion. The PBGC insures only about two-thirds of private sector single-employer-defined benefit plans, but almost all multi-employer plans (Pension Benefit Guarantee Corporation 2009, p. 5). Among the types of defined benefit plans PBGC does not insure are small (fewer than 25 employees) plans maintained by small professional service employers like doctors, lawyers, and accountants. Since the PBGC excludes only the smaller single-employer plans, and includes most multi-employer plans, we assume that it covers well more than 66 percent (i.e., two-thirds) of all liabilities, setting our estimate at 90 percent.

Other Financial Firms—Implicitly Guaranteed Liabilities:

"Total liabilities of AIG" found in its 10K report, less "estimated insured deposits" of AIG Federal Savings Bank found on the FDIC's website in the "Institution Directory" (http://www2.fdic.gov/idasp).

Other Financial Firms—Total Liabilities:

Board of Governors (2010), Tables L.116—Property-Casualty Insurance Companies; L.117—Life Insurance Companies; L.126—Issuers of Asset-Backed Securities; L.127—Finance Companies; L.128—Real Estate Investment Trusts; L.129—Security Brokers and Dealers; L.131—Funding Corporations, less taxes payable whenever a figure for taxes was reported on these tables.

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Economic Quarterly—Volume 96, Number 3—Third Quarter 2010—Pages 291–317

The Politics of Sovereign Defaults

Juan Carlos Hatchondo and Leonardo Martinez

Sovereign debt issuance and repayment decisions are determined by public officials and may thus be affected by issues such as the proximity of elections; conflicts between the executive branch and the parliament; institutional breakdowns such as military coups; etc. This article first discusses theoretical and empirical studies about the role of political factors in sovereign default episodes. Before concluding, the article also discusses the role of political factors in five recent default episodes.[1]

The preferences of public officials and the environment in which they must act affect their perceived costs and benefits of defaulting. This has been recognized by several authors. For instance, in discussing the role of political factors as determinants of defaults, Sturzenegger and Zettelmeyer (2006) conclude that "a solvency crisis could be triggered by a shift in the parameters that govern the country's willingness to make sacrifices in order to repay, because of changes in the domestic political economy (a revolution, a coup, an election, etc.)..." Similarly, Rijckeghem and Weder (2009) argue that a country's willingness to pay is influenced by politics, i.e., by the distribution of political power and of benefits and costs of defaulting across voters. The heterogeneity of public officials' preferences is also highlighted by Santiso (2003) who writes, "One basic rule of the confidence game [in international financial markets] is then to be very careful when nominating the official government voicer. For investors it is mainly the ministry of economics or finance or the governor of the central bank."

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[1] Hatchondo, Martinez, and Sapriza (2007a) present a brief discussion of political costs of defaulting. In this article, we extend their analysis and present a more thorough discussion of theoretical and empirical results.

We first describe theoretical studies that illustrate how the risk of losing elections may induce a sovereign to avoid a default even when creditors have no access to legal procedures that would allow them to force the sovereign to pay. This risk would be present when sovereign debt is at least partially held by local creditors with political power to deny support to political groups that advocate for a default. Note that, since it is difficult to declare a selective default on foreign bondholders only, the presence of these local creditors could also explain why foreign investors are willing to buy sovereign debt.

Second, we describe theoretical work that studies how political turnover, i.e., the alternation in office of policymakers with different objectives, affects incentives to borrow from foreign lenders and to default on debt held by foreigners. Policymakers may differ in the weights they assign to different constituencies of domestic residents when allocating fiscal resources and they may differ in their willingness to pay the debt. Studies that assume differences in policymakers' spending preferences find that a higher frequency of political turnover tends to generate higher debt levels and higher default probabilities. In contrast, studies that assume that policymakers differ in their willingness to repay debt find that the relationship between the default probability and the frequency of political turnover may be nonmonotonic.

Studies that assume that policymakers differ in their willingness to repay make possible the existence of defaults triggered by political turnover.[2] We refer to such default episodes as "political defaults." Political defaults occur when a "creditor-friendly" government (with a higher willingness to pay) is replaced by a "debtor-friendly" government (with a lower willingness to pay). It should be mentioned that while political turnover may explain the timing of the default decision, poor economic conditions are likely to play a key role in political defaults. In fact, in Hatchondo, Martinez, and Sapriza's (2009) model of political defaults, political defaults are only likely to occur after a creditor-friendly government encounters poor economic conditions that lead it to choose high borrowing levels. These studies also find that after political defaults, debt and interest rate spread levels are lower than the levels observed after defaults caused by negative income shocks only, and are lower than the pre-default levels.[3] Recall that a political default is triggered when a creditor-friendly government is replaced by a debtor-friendly government. These studies argue that in a political default, post-default debt levels are

[2] Using a historical data set with 169 sovereign default episodes, Tomz and Wright (2007) find that 38 percent of default episodes in their sample occurred in years when the output level in the defaulting country was above the trend value. Thus, it is unlikely that these episodes were triggered by difficult economic conditions. Tomz and Wright argue that some of these episodes may have been triggered by political turnover.

[3] The interest rate spread corresponds to the difference between the yield of sovereign bonds and the risk-free rate. When we contrast theoretical predictions with data, we use the yield on 90-day U.S. Treasury bills as the risk-free rate.

lower than pre-default levels because investors are less willing to lend to debtor-friendly governments. This contrasts with alternative explanations that rely on a boycott against a government in default that is not explained by characteristics of the government but by its past behavior.[4]

Third, we review empirical studies that have tested the existence of statistical relationships between political factors and default decisions. These studies have found that the proximity of elections, the turnover of government officials, increases in political instability, and the presence of a presidential democratic regime instead of a parliamentary democratic regime are statistically associated with a higher default probability.

We conclude with a brief description of the role of political turnover in five recent default episodes: Argentina 2001, Ecuador 1998, Pakistan 1998, Russia 1998, and Uruguay 2003. First, we attempt to identify whether these default episodes occur after a creditor-friendly government was replaced by a debtor-friendly government. In order to do so, we look at a measure of political risk computed by the International Country Risk Guide (ICRG). We argue that if a creditor-friendly government was replaced by a debtor-friendly government at the time of the default, the level of political risk computed by the ICRG should be lower in the years before the default than in the years after the default. We find that only in Argentina is the level of political risk systematically lower in the years before the default than in the years after the default. We also present anecdotal evidence indicating that political turnover was important in determining the timing of the Argentine default. The role of political factors in the Argentine crisis has also been highlighted in previous studies. IMF (2004) argues that in Argentina "economic, social, and political dislocation occurred simultaneously, leading to the resignation of the President, default on Argentina's sovereign debt, and the abandonment of convertibility..." Similarly, IMF (2003) finds that in Argentina "the inability to mount a policy response stemmed from a combination of economic constraints and political factors..." In addition, we show that the behavior of interest rate and debt levels before and after the Argentine default is broadly aligned with the predictions of theoretical studies.

1. THEORETICAL LITERATURE

In this section, we summarize lessons that can be extracted from theoretical studies that analyze the role of political factors in sovereign default episodes.

[4] For instance, it is often argued that creditors may punish a defaulting government by excluding it from capital markets. This is assumed in Eaton and Gersovitz's (1981) seminal model of sovereign default and in extensions of their work (Hatchondo, Martinez, and Sapriza [2007b] discuss the role of this assumption).

First, we discuss political costs of sovereign defaults. Second, we describe how political turnover affects debt issuance and repayment decisions.

Political Costs of Defaults

In a hypothetical scenario in which sovereign defaults were costless, governments would always default and, in anticipation of that, investors would not purchase public debt to begin with. Yet, we observe that governments are able to borrow significant amounts in spite of the weak legal protection enjoyed by bondholders. This observation can be taken as evidence of costs associated with sovereign defaults.[5] The literature has debated the ability of foreign creditors to impose explicit sanctions on governments that have reneged on their debts (see Hatchondo, Martinez, and Sapriza [2007a]). This section reviews a number of studies that emphasize that sovereign defaults may be politically costly primarily because a fraction of sovereign debt is held by local voters.

For a government, an alternative to defaulting is to raise taxes in order to be able to pay its debt. In any society, people are likely to have different exposures to the debt of their government and to a tax increase. In general, we can expect that a sovereign default will not occur as long as debtholders have sufficient political power. Dixit and Londregan (2000) formalize this idea. They argue that when making the decision to raise taxes to pay the interest or repay the principal on its debt, the government will pay due attention to the relative political power of the debtholders and other taxpayers. They consider a two-period model in which debt is issued in the first period and two political parties compete to win an election. Voters differ in their learning abilities for human capital accumulation, initial wealth, and in their preferences over "position issues" such as gun control, abortion, etc. In the first period, voters invest by accumulating human capital or by buying government bonds. Government debt revenues are allocated to build infrastructure capital. Elections are held at the end of the first period. Before the elections, each party presents a platform of income taxes, debt repayment, and their stance on position issues. In the second period, production takes place and the party in office levies taxes and decides the fraction of debt that is repaid. There are no punishments to a defaulting government. Dixit and Londregan (2000) show that under some distributional assumptions, the number of bondholders who are indifferent to voting for any of the two parties on the basis of position issues alone may be larger than the number of nonbondholders—voters that decided to invest in human capital instead of buying government bonds. Consequently, an

[5] That is, for sovereign debt to exist, it is necessary that at least in some circumstances it would be more costly for a sovereign to default than to pay back its debt. Similarly, for sovereign defaults to exist, it is necessary that at least in some circumstances it would be more costly for a sovereign to pay back its debt than to default.

equilibrium with positive debt issuance and no default can be sustained. In that equilibrium, the presence of a larger number of swing voters who would favor the repayment of debt ensures that the party that proposes to pay back the debt wins the election.

In general, citizens who are wealthier may hold more government debt and may suffer more in the event of a sovereign default. Furthermore, those who are older tend to be wealthier while those who are younger tend to generate more income and thus are more exposed to an increase in income taxes. Consequently, as long as wealthier and older citizens impose their will, sovereign defaults may be prevented. Tabellini (1991) emphasizes these ideas. He presents a two-period economy that is inhabited in period 1 by a generation of young agents who live for two periods. There are two generations active in period 2: young and old. The young in period 1 differ in their initial endowment of goods. The old in period 2 care about the welfare of their offspring, so they may leave bequests. There is political competition and the government's decisions are determined by majority voting. Debt is issued in the first period and the repayment decision is made in the second period. Tabellini (1991) assumes that individuals cannot punish a defaulting government. The young in period 1 vote on how much debt to issue and individually decide how much to save for the next period. They can only save in government bonds, so aggregate savings must equal total debt issuance (there is no external debt). At the beginning of period 2, the young and old vote on how much debt is going to be repaid. Old agents are the debtholders and young agents are the only ones being taxed in period 2. Tabellini (1991) shows that a coalition composed of old agents and those young agents who are the children of the wealthy (old) bondholders may vote to repay the debt. The second group in the coalition may enjoy a net benefit from repaying the debt because the taxes they pay to honor the debt are offset by the bequests they plan to receive from their parents.

If a sovereign cannot perfectly discriminate between domestic and foreign creditors, then the political cost of defaulting described above also allows the sovereign to issue debt to foreign lenders. This is argued by Guembel and Sussman (2009). While the setups presented in Tabellini (1991) and Dixit and Londregan (2000) do not consider the possibility of a government borrowing from foreign lenders, Guembel and Sussman (2009) study this possibility. They propose an environment in which the government cannot perfectly discriminate between domestic and foreign creditors who cannot impose sanctions to a government in default. The authors assume electoral competition between two political parties that compete with a platform that specifies the repayment strategy to be implemented once in power. The two parties are identical and their only objective is to win elections. A default entails a redistribution of wealth toward local individuals who hold an amount of debt that is low enough to make the loss caused by not being compensated for the defaulted debt smaller than the benefit from avoiding the taxes they would have paid to service the debt. Like Tabellini (1991) and Dixit and Londregan (2000), Guembel and Sussman (2009) show that, under certain circumstances, the government repays its debt in spite of the fact that no creditor can punish defaulting governments. The reason is that the median voter would favor a platform that proposes to pay back the debt. The authors also show that when there is imperfect information about the characteristics of the median voter, "lending booms" (high foreign demand for bonds) can price the median voter out of the market and thus increase the probability of default. With imperfect information, investors may mistakenly interpret high bond prices that can be caused by an increase in foreign demand for bonds as evidence of a strong willingness to pay by the future government.

Drazen (1998) focuses on analyzing the influence of political factors on a government's decision to finance its expenditures by issuing debt to domestic or foreign lenders. Like Guembel and Sussman (2009), Drazen (1998) considers a setup in which the government issues debt to local and foreign residents. Unlike Guembel and Sussman (2009), Drazen (1998) studies the case in which the government can selectively default on local or foreign debt and foreign debtholders can punish a defaulting government. He argues that governments can, in fact, exert some control over whether debt is held by domestic or foreign residents. In particular, he mentions that the government can affect the allocation of debt among domestic and foreign agents through capital controls that restrict the ability of domestic (foreign) residents to buy debt issued abroad (locally), through the currency denomination of public debt (debt denominated in domestic currency may be more attractive to domestic residents), through differential tax treatment, etc. He proposes then a political economy model in which domestic debtholders vote on repayment decisions. Thus, as in Tabellini (1991) and Dixit and Londregan (2000), debt held by local agents can be sustained in equilibrium. Since Drazen (1998) assumes that foreign debtholders *can* punish a defaulting government, debt held by foreigners can also be sustained in equilibrium. Drazen (1998) argues that countries where debtholders have more political power should tend to finance a higher proportion of public expenditures by issuing domestic debt. In his setup, as the income distribution becomes less concentrated, more agents can save and buy domestic debt and thus benefit from interest payments on public debt. Those agents would vote for a political platform that proposes to issue more domestic debt and to honor this debt. Consequently, countries with relatively less concentrated income distributions (higher median income for the same mean) may tend to finance a higher proportion of public expenditures by issuing domestic debt.

In summary, the main lessons from the literature on the political costs of sovereign defaults described above are (i) a sovereign default will not occur as long as local debtholders have sufficient political power; (ii) a default is likely to be prevented as long as wealthier and older citizens impose their will; (iii) as long as a sovereign cannot default solely on foreign creditors without affecting local creditors, political costs of defaulting also allow the sovereign to issue debt to foreign lenders; and (iv) countries where debtholders have more political power (for instance, because of a more even distribution of income) will tend to finance a higher proportion of public expenditures by issuing domestic debt.

Political Turnover and Sovereign Default Risk

In this section, we summarize lessons from studies that focus on the role of political turnover, which is defined as the alternation in power of groups with different preferences. These studies typically lack a deep theory that links the objectives of citizens and policymakers, and links policy choices to election outcomes. This modeling strategy may be useful to clarify causality relationships from political variables to sovereign debt issuance and default decisions.

An economy is said to have more political stability when political turnover is less frequent. What is the relationship between political stability and default risk? Amador (2003) and Cuadra and Sapriza (2008) contribute to answering this question. They study models of sovereign default in which policymakers disagree on the optimal allocation of fiscal resources within each period because they want to please different constituencies. They show that an increase in political stability reduces the risk of a sovereign default, which in turn reduces the interest rate spread on sovereign bonds. The intuition for their results is as follows. The current government knows that future resource allocations may be decided by a government that would make different choices from the ones the current government would make. Consequently, the current government would like to transfer resources from the future (when decisions may not be made following its preferred criteria) to the present (when it can decide where to allocate those resources). With less political stability it is more likely that the current government will disagree with the choices of future governments and, therefore, the current government is more eager to transfer resources from the future to the present. Thus, political stability affects the effective discount factor of the incumbent government. One instrument that the government can use to bring resources from the future is to issue more debt. Higher debt levels increase the default probability—when defaulting, the government benefits from not paying back its debt and these benefits are larger if debt levels are larger. Another strategy is defaulting in situations in which the government would have to pay large amounts in the present while a substantial fraction of the cost of defaulting appears in the future. In contrast, if there is more political stability, the government is less eager to transfer resources to the present, it wants to borrow less, and it is less willing to default.

The analysis in Amador (2003) and Cuadra and Sapriza (2008) assumes that policymakers do not differ in their willingness to pay back sovereign debt and, therefore, receive the same treatment from international investors. This implies that lenders do not care about the type of policymaker in office or about which type of policymaker may be in office in the future. This seems unrealistic. Examples abound in which politicians disagree about the benefits of maintaining a good credit standing. As explained in the next section, in the proximity of elections, default risk may be influenced by poll data. This suggests that a better understanding of the relationship between political turnover and default risk could be achieved by allowing for the existence of policymakers with different preferences for debt repayment.

Aghion and Bolton (1990) study a setup in which policymakers differ in their willingness to pay. Unlike other studies described in this section, they present a model with endogenous turnover. They show how the government may want to overaccumulate debt to affect the result of elections. They consider a two-period closed economy inhabited by a continuum of agents who live for two periods (there are no intergenerational transfers). At the beginning of each period, elections are held to appoint government authorities. Agents only differ in the endowment they receive in every period and derive utility from private consumption and a publicly provided good. The first-period government determines the level of the public good provided in that period and the proportion of expenditures that is financed through a uniform tax and through debt issuances. The second-period government determines the level of the public good provided in that period, the uniform tax, and the repayment of debt. Aghion and Bolton (1990) assume that there are two political parties. The "right-wing" ("left-wing") party is assumed to maximize the utility of a group of agents with an above-average (below-average) income level. Given that debtholdings increase with income, the right-wing party displays a stronger preference to pay back debt than the left-wing party. By issuing more debt in the first period, the right-wing party increases the size of the constituency that prefers the debt to be paid back and, through that, it increases the likelihood of winning the election held at the beginning of the second period. Thus, electoral concerns induce the right-wing party to issue a larger amount of debt in the first period.

Cole, Dow, and English (1995), Alfaro and Kanczuk (2005), and Hatchondo, Martinez, and Sapriza (2009) also study models of sovereign default with two types of policymakers that differ in their willingness to pay. Unlike Aghion and Bolton (1990), these studies assume that the two types alternate stochastically in power. In their setups, policymakers who assign more weight to the future (for example, because they are more likely to win elections) are more willing to pay because they are more concerned about the costs of defaulting that appear in the future.

Cole, Dow, and English (1995) and Alfaro and Kanczuk (2005) study setups with asymmetric information about the type of policymaker in office. Thus, a cost of defaulting is that lenders update their beliefs about the government's type, which in turn may affect future borrowing opportunities. Cole, Dow, and English (1995) show that an equilibrium exists in which the patient policymaker always repays, the impatient policymaker always defaults, and, in the period where there is a type change from impatient to patient, the patient policymaker is able to perfectly signal its type by making a settlement payment because the impatient type would not find it optimal to do the same. Their model can explain cycles of borrowing and exclusion from credit markets that finish when the government pays part of the debt in default. They argue that this pattern is consistent with the aftermath of many 19th century default episodes in Latin America and in the United States.

In the framework proposed by Alfaro and Kanczuk (2005), there are equilibria in which lenders do not know the type of policymaker in office. They allow for a publicly observable aggregate productivity shock and show the existence of equilibria in which, for moderately negative productivity shocks, the patient type does not default in order to avoid damaging the government's reputation—i.e., the probability that lenders assign to the patient type being in office.

In order to simplify the learning process faced by lenders and make their models tractable, Cole, Dow, and English (1995) and Alfaro and Kanczuk (2005) limit the set of borrowing levels available to the government. In general, in models with asymmetric information, equilibrium borrowing levels may be distorted by the desire of the borrower who is less willing to default to reveal his type through his borrowing choice. In particular, when borrowing less would allow a patient government to distinguish itself from impatient governments, the patient government may not want to borrow as much as it would if its type was public information. The drawback of restricting the set of borrowing levels is that it limits the usefulness of the models for studying macroeconomic fluctuations.

Hatchondo, Martinez, and Sapriza (2009) consider a political process similar to the one used by Cole, Dow, and English (1995) and Alfaro and Kanczuk (2005), but do not assume asymmetric information about the government type and, therefore, do not need to restrict the set of borrowing levels available to the government. Moreover, the framework used by Hatchondo, Martinez, and Sapriza (2009) follows closely the one used in recent quantitative models of sovereign default (see, for example, Aguiar and Gopinath [2006], Arellano [2008], Hatchondo and Martinez [2009], and Hatchondo, Martinez, and Sapriza [2010]).

Hatchondo, Martinez, and Sapriza (2009) identify two channels through which political stability may influence default risk in addition to the channel outlined in Amador (2003) and Cuadra and Sapriza (2008). On the one hand, if political turnover is expected to trigger a default, the default risk premium charged on bond issuances is higher when the probability of political turnover is higher (which corresponds to lower political stability). Thus, this channel predicts a negative relationship between political stability and default risk, as in Amador (2003) and Cuadra and Sapriza (2008). On the other hand, if less political stability were to imply more default risk, the default risk premium charged on bond issuances would be higher. In turn, a higher borrowing cost would make the government less willing to borrow. In particular, it could make the government unwilling to choose debt levels for which a political default—defined as a default that would occur because of political turnover—would be likely. Therefore, less political stability could reduce default risk. The possibility of a positive relationship between political stability and default risk is not present in Amador (2003) and Cuadra and Sapriza (2008).

Based on their findings on the relationship between political stability and borrowing costs, Hatchondo, Martinez, and Sapriza (2009) argue that political defaults are only likely to occur in economies where there is enough political stability. If the current government chooses borrowing levels that would lead to a default after political turnover, it has to compensate lenders for this contingency, i.e., for the contingency of another government becoming the decisionmaker in the future. If the probability of this contingency is high enough (political stability is low), it is too expensive for the current government to choose borrowing levels that would lead to a political default. In this scenario, the current government does not borrow so heavily and, therefore, political turnover would not trigger a default.

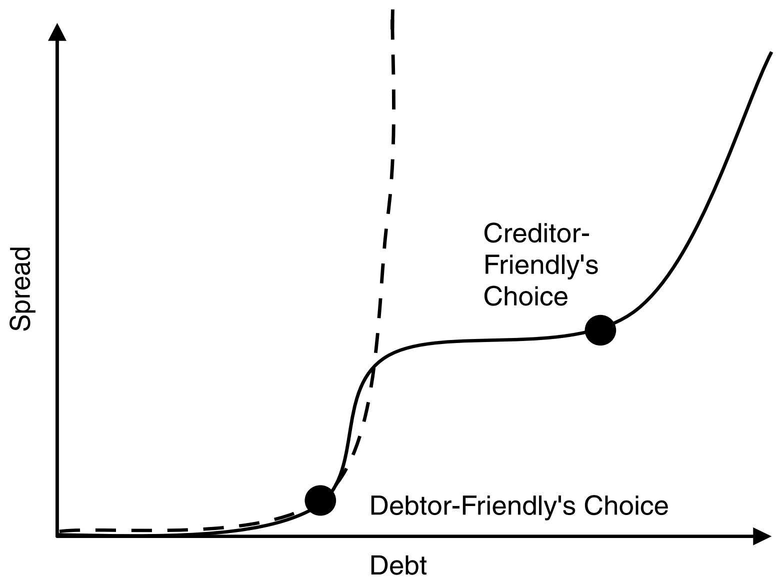
In addition, Hatchondo, Martinez, and Sapriza (2009) show that, in economies with enough political stability, political turnover may weaken the correlation between default and output. Thus, introducing political turnover may bring the predictions of the baseline quantitative model of sovereign default closer to the data. Using a historical data set with 169 sovereign default episodes, Tomz and Wright (2007) report a weak correlation between economic conditions and default decisions. They find that 38 percent of default episodes in their sample occurred in years when the output level in the defaulting country was above the trend value.

The model presented by Hatchondo, Martinez, and Sapriza (2009) also highlights distinctive features of political defaults. In their model, if a default is not preceded by political turnover, post-default debt levels tend to return to pre-default levels relatively fast. In contrast, if a default is caused by political turnover, post-default debt levels tend to be lower than pre-default levels. Recall that a default is caused by political turnover when a government is replaced by another government that is more willing to default. In a political default, post-default debt levels are lower than pre-default levels because the cost of borrowing is higher for governments that are more willing to default and, consequently, post-default governments borrow less than pre-default governments. This contrasts with alternative explanations for low post-default borrowing levels that rely on a boycott against a government in default that is not explained by characteristics of the government but by its previous default decision (for instance, creditors could agree to exclude a defaulting government from capital markets independently of the likelihood of future government repayments). The mechanism that generates lower post-default debt levels illustrated by Hatchondo, Martinez, and Sapriza (2009) is similar to one presented by Cole, Dow, and English (1995). In Cole, Dow, and English (1995), post-default governments cannot borrow because they would always default. In Hatchondo, Martinez, and Sapriza (2009), post-default governments can borrow but at a higher interest rate than pre-default governments. In equilibrium, post-default governments *choose* to borrow less than pre-default governments.

The second distinctive feature of political defaults highlighted by Hatchondo, Martinez, and Sapriza (2009) is that post-default equilibrium spreads tend to be lower than pre-default spreads. That is, high-willingness-to-pay governments pay higher spreads than do low-willingness-to-pay governments. Before a political default, when the government has a high willingness to pay, bondholders require a compensation for the possibility that the current government is replaced by a government with a lower willingness to pay. In contrast, after a political default, the low-willingness-to-pay government does not need to compensate lenders for the risk of political turnover. This is because political turnover would actually mean good news to bondholders.

To further illustrate the relationship between pre- and post-political default levels of debt and spread, consider the case in which two types of governments, creditor-friendly and debtor-friendly, alternate in power, and a political default occurs when the first type is replaced by the second type. In addition, suppose that the type of policymaker currently in charge of the government is likely to be in charge of the government at the time the debt it issues has to be paid back (but a change in the type of policymaker is possible). Figure 1 presents the interest rate spread each of these two types of government would have to pay as a function of the borrowing level they choose. The functions in the figure resemble the equilibrium functions derived in Hatchondo, Martinez, and Sapriza (2009). The functions presented in Figure 1 display three steps. The first step corresponds to "low" issuance volumes. At these volumes, the debt issued is sufficiently low that the government will almost surely pay it back, regardless of the type in power. The second step corresponds to "intermediate" issuance levels. These are the issuance values such that a debtor-friendly policymaker would default in the next period whereas a creditor-friendly policymaker would pay. When a creditor-friendly policymaker is in office, the spread charged by lenders for these issuance volumes is increasing in the probability of political turnover. When a debtor-friendly policymaker is in office, the spread charged by lenders for these issuance

Figure 1 Interest Rate Spread



volumes goes to infinity because a debtor-friendly government would choose to default on these volumes (this is the case when the recovery rate on defaulted debt is zero, as in Hatchondo, Martinez, and Sapriza [2009]). Finally, the third step corresponds to "high" issuance volumes. At these volumes, investors realize that the government will almost surely default tomorrow, regardless of the type in power and, therefore, spreads go to infinity. Hatchondo, Martinez, and Sapriza (2009) show that, when facing such options, creditor-friendly governments may choose to issue intermediate debt levels and to pay intermediate spreads while debtor-friendly governments may choose to issue low debt levels and to pay low spreads. Thus, the levels of debt and spread are typically higher before a political default than after the default. Figure 1 also presents the typical government's choices according to the equilibrium studied by Hatchondo, Martinez, and Sapriza (2009).

In summary, the literature studying the relationship between political turnover and default risk shows us that: (i) governments may want to over-accumulate debt to affect the result of elections; (ii) more political stability may imply a lower default risk if it makes the government less eager to transfer resources to the present; (iii) political defaults are only likely to occur in economies where there is enough political stability; (iv) political turnover may weaken the correlation between default and output; (v) around political defaults, post-default debt levels may be lower than pre-default levels; and (vi) creditor-friendly governments may pay higher spreads than debtor-friendly governments and, consequently, post-political-default spreads may be lower than pre-political-default spreads.

2. EMPIRICAL LITERATURE

In Section 1, we discussed insights from theoretical studies that show how political factors may influence sovereign default risk. In this section, we summarize the findings of empirical studies that have investigated statistical relationships between political factors and default risk. These studies have found that the proximity of elections, changes in the finance minister or central bank governors, increases in indicators of political instability, and the presence of a presidential democratic regime instead of a parliamentary democratic regime are statistically associated with a higher default probability. These studies include controls such as the debt over gross domestic product ratio, the level of reserves, or output growth. This attenuates the criticism that political indicators may be significant only because of their correlation with policy choices (such as the accumulation of debt).

Political Stability

In Section 1, we discussed how an increase in political instability may increase default risk. We discuss next studies that propose measures of political stability and use these measures to evaluate whether political stability affects default risk. Citron and Nickelsburg (1987), Balkan (1992), and Brewer and Rivoli (1990) find that this seems to be the case.

Citron and Nickelsburg (1987) use a logit model to estimate the probability of default using data from Argentina, Brazil, Mexico, Spain, and Sweden for the 1960–1983 period. They construct an indicator of political instability that measures the number of changes in government—that were accompanied by changes in policy—that took place within the previous five years. They find that, on top of various macroeconomic indicators, their measure of political instability has a significantly positive effect on the default probability.

The results in Balkan (1992) are consistent with the ones in Citron and Nickelsburg (1987). Balkan (1992) uses an index of political instability that "measures the amount of social unrest that occurred in a given year." He estimates the probability of default using a sample larger than the one used by Citron and Nickelsburg (1987): 31 countries from 1971–1984. Controlling for 10 economic indicators and an index of democratization, he finds that a higher index of political instability increases the probability of observing a debt rescheduling in the subsequent year.

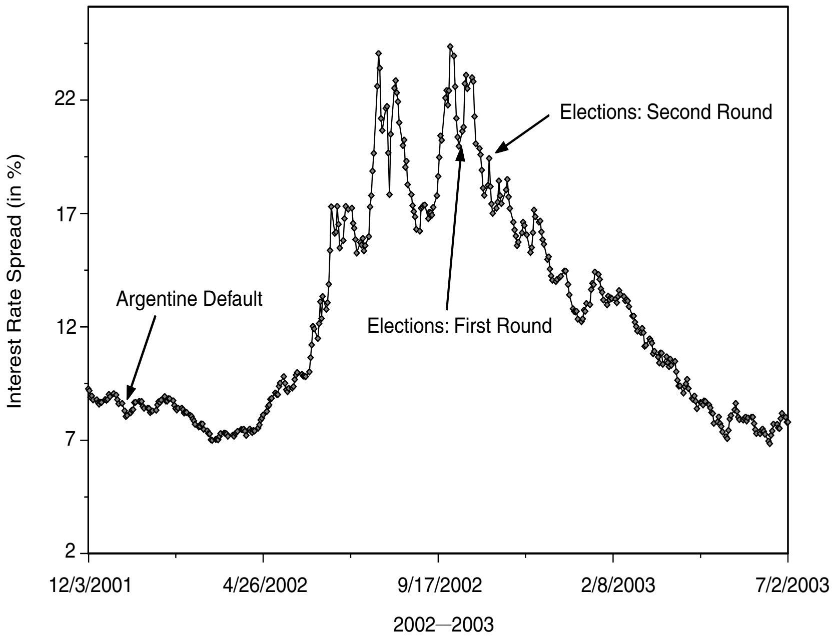
Brewer and Rivoli (1990) also find that political instability has a significant negative effect on a country's perceived creditworthiness. In particular, they argue that the frequency of regime change appears to be at least as important as economic variables in explaining lenders' risk perceptions. They use two indexes of regime stability. One index represents the frequency of change in the head of government and the other represents the frequency of change in the governing group (political party or military government). Instead of using data on defaults or interest rate spread, Brewer and Rivoli (1990) use credit ratings from Institutional Investor and Euromoney (two private credit-rating consultants). They use a sample of 30 countries from 1967–1986. They do not find evidence in favor of other political indicators such as the existence of an armed conflict or the democratic nature of the government having significant effects on credit ratings.

Bussiere and Mulder (2000) use a sample of 44 developing countries to test the contribution of political variables to the severity of the financial crises that took place between 1994–1997 (not all crises are linked to a default episode). They find that indicators about the uncertainty of election outcomes amplify the magnitudes of subsequent crises. Those indicators consist of an index of volatility of the electorate (the change in the proportion of seats held by each party from one election to the other) and a dummy variable that captures the presence of elections during the sample period. (They also find that an index of political polarization based on the number of political parties and an index of the fragility of the ruling coalition do not have statistically significant effects.)

Political Turnover and Default Risk

In Section 1 we also discussed theoretical studies that assume that policymakers differ in their willingness to default, which allows for political defaults—i.e., defaults triggered by political turnover—to occur. Figure 2 illustrates a notable example of how the probability of default (reflected in sovereign bond spreads) may be influenced by changes in the probability of political turnover. This should happen when policymakers differ in their willingness to default. The figure shows the behavior of the sovereign spread in Brazil before and after the election of 2002. The concerns raised by the possible electoral victory of the left-wing candidate Luiz Inacio "Lula" Da Silva because of his previous declarations in favor of a debt repudiation is the most accepted explanation for the sharp increase in the spread on sovereign bonds preceding the 2002 Brazilian election. Goretti (2005) finds further evidence in favor of that hypothesis. She uses a nonlinear econometric model to account for the behavior of the sovereign spread in Brazil between November 2001 and October 2002. She finds that a measure of the perceived probability of Lula's victory (based on opinion polls) has a statistically significant effect on spread levels. In the event, Brazil did not default on its debt.

Figure 2 Elections and Sovereign Bond Spread in Brazil



Source: J.P. Morgan (EMBI Global).

Notes: The Emerging Market Bond Indices (EMBI) track the return on traded debt instruments nominated in a foreign currency. J.P. Morgan computes various indices that differ in the countries included, the weights assigned to countries, and the liquidity of the debt instruments included.

The results in Block and Vaaler (2004) and Manasse, Roubini, and Schimmelpfennig (2003) suggest that the Brazilian example illustrated in Figure 2 is not an exception. Close to elections, the possibility of political turnover seems to increase the level of default risk. Block and Vaaler (2004) find that election years are associated with an average downgrade of sovereign debt. They also report that bond spreads are higher in the 60 days before an election compared to spreads in the 60 days after an election. They study a sample of 19 developing countries from 1987–1998. The sample includes 18 presidential elections. Similarly, Manasse, Roubini, and Schimmelpfennig (2003) find that the probability of a debt crisis increases in years with presidential elections. They define a debt crisis as either an episode classified as a default by Standard & Poor's, or the acceptance of an IMF loan in excess of 100 percent of the country's quota. They use a sample of 37 developing countries from 1976–2001 and estimate the probability of a debt crisis one year ahead.

The equilibrium behavior predicted by Hatchondo, Martinez, and Sapriza (2009) may help us understand why an increase in the probability of political turnover, on average, increases default risk, as found by Block and Vaaler (2004) and Manasse, Roubini, and Schimmelpfennig (2003). Hatchondo, Martinez, and Sapriza (2009) show that the effect of political turnover on the default probability may depend on the type of the *current* government. In their model, when a debtor-friendly government is in office, the level of default risk does not depend on the probability of political turnover because political turnover would not trigger a political default. In contrast, when a creditor-friendly government is in office, the level of default risk increases with respect to the probability of political turnover because political turnover could trigger a political default. Thus, on average, one can expect that the possibility of political turnover close to elections would increase the level of default risk, as found in empirical studies.

It should be stressed that a change in the type of government in power does not need to be preceded by an election. For instance, the turnover of high rank government officials could signal changes in a government's willingness to default. Moser (2007) and Moser and Dreher (2007) find evidence suggesting that this may be the case. Moser (2007) finds that changes in the finance minister generate an average increase of 100 basis points of the sovereign spread on the day of the announcement. This is based on a sample of 12 Latin American countries from 1992–2007. Moser (2007) documents that around one third of the announcements of a change in the finance minister during that time led to a decrease in the sovereign spread, which implies that the increase in the spread of the negative announcements is larger than 100 basis points. Similarly, based on a sample of 20 emerging countries from 1992–2006, Moser and Dreher (2007) find that bond spreads increase and local currencies depreciate as a result of changes in central bank governors.[6]

As discussed in Section 1, policymakers may differ in their willingness to default because they represent constituencies with different exposures to sovereign debt. The political power of debtholders may vary with the characteristics of the political system. Consequently, these characteristics could affect default decisions. The findings in Saiegh (2009), Kohlscheen (2009), and Rijckeghem and Weder (2009) suggest that this is the case.

Using a sample of 48 developing countries between 1971–1997, Saiegh (2009) finds that countries governed by a coalition of parties are less likely to default than those governed by single-party governments. Similarly, Kohlscheen (2009) finds that parliamentary democracies display a lower

[6] A possible caveat of these results is that the political factors may reflect shocks to fundamentals.

probability of default compared to that of presidential democracies. He estimates a probit model based on a sample covering 59 democracies from 1976–2003.

Rijckeghem and Weder (2009) classify regimes as democratic and nondemocratic according to the value of a democratization index, and differentiate between defaults on external and domestic debt. They use a sample of 73 countries from 1974–2000. Rijckeghem and Weder (2009) find that the frequency of defaults on external debt is larger than the frequency of defaults on domestic debt, independent of whether the political regime is democratic. When they restrict their estimations to samples with only democratic regimes, they find that parliamentary systems and systems with a large number of veto players deter external defaults as long as economic conditions are sufficiently good. They do not find statistical evidence of other political factors deterring defaults on domestic debt. For nondemocratic regimes, they find that the proximity to elections and low polarization do deter defaults on domestic debt but they do not find evidence that political indicators other than the type of regime deter defaults on external debt.

Balkan (1992) constructs an index of democracy that "measures the extent that the executive and legislative branches of government reflect the popular will." He estimates the probability of default using a sample of 31 countries from 1971–1984. Controlling for 10 economic indicators and a measure of political stability, he finds that a higher index of democracy decreases the probability of observing a debt rescheduling in the subsequent year.

3. RECENT SOVEREIGN DEFAULTS IN EMERGING MARKETS

In this section, we discuss the influence of political factors on five recent default episodes: Argentina 2001, Ecuador 1999, Pakistan 1999, Russia 1998, and Uruguay 2003. First, we attempt to identify whether these defaults were political defaults. We do this with a commonly used index of political risk. This index suggests that the Argentine default is the most likely to have been political. Then, we present the behavior of the levels of sovereign debt and spreads around the Argentine default and show that the Argentine data is consistent with the predictions of the theory developed by Hatchondo, Martinez, and Sapriza (2009) for political defaults.

Political Turnover and the International Country Risk Guide Aggregate Index of Political Risk

In Section 1, we explain how governments may differ in their willingness to pay back sovereign debt because they represent different constituencies. For instance, while some governments may be more concerned about the well-being of debtholders, others are more concerned about the well-being of taxpayers. We also explain that this implies that a default may occur when a creditor-friendly government (with a lower willingness to default) is replaced by a debtor-friendly government (with a higher willingness to default) and we refer to such a default as a political default. Having a measure of governments' willingness to default would allow us to conduct a systematic analysis of whether default episodes were triggered by political turnover. We will use as such a measure the index of political risk for investors included in the International Country Risk Guide (ICRG). ICRG is a credit-rating publication published by The Political Risk Services Group. This index is commonly used in empirical studies (see, for example, Erb, Harvey, and Viskanta [1996, 1999], Bilson, Brailsford, and Hooper [2002], Reinhart, Rogoff, and Savastano [2003], and Bekaert, Harvey, and Lundblad [2007]).

Bilson, Brailsford, and Hooper (2002) define political risk as "the risk that arises from the potential actions of governments and other influential domestic forces, which threaten expected returns on investment." In the context of sovereign debt, default is the government's action that affects the return obtained by lenders and, for a given debt level, political risk for investors is lower (higher) when policymakers with a high (low) willingness to pay are in power. Thus, political turnover could trigger a default when the level of political risk changes from low to high.[7]

The ICRG index of political risk is one of the three components of the overall ICRG country risk index. The other two indexes are the financial risk index and the economic risk index. The index of political risk is supposed to reflect political risk only, independent from economic risk and financial risk (which are captured by the other two indexes). Thus, the index of political risk does not necessarily mirror default risk. In fact, we will illustrate that the default premium implied by Argentine bond prices (the spread) was higher when political risk was lower.

The ICRG Index of Political Risk and Political Turnover in Recent Default Episodes

Table 1 presents summary statistics of the behavior of political risk (100 minus the ICRG index of political risk) before and after the default episodes in Argentina 2001, Ecuador 1999, Pakistan 1999, Russia 1998, and Uruguay 2003. Since a political default occurs after a creditor-friendly government is replaced by a debtor-friendly government, one should expect that in the years before a political default political risk was lower than in the years after

[7] It must be said that the ICRG index of political risk does not purely reflect an assessment about the type of policymakers in office. It also depends on the perceived likelihood of observing a change of the type in office and on institutional factors.

Table 1 Political Risk in Recent Default Episodes

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Argentina	12-2001	29.1	25.6	38.4	0	0	124	n/a
Ecuador	07-1999	40.3	39.3	42.9	11.5	0	28	n/a
Pakistan	11-1999	49.7	48.6	52.9	27.1	8.3	4	18
Russia	08-1998	44.6	43.3	47.5	34.2	22.2	22	23
Uruguay	05-2003	27.8	27.9	27.5	36.5	61.1	1	13

Notes: (1) month of default; (2) average risk in the sample; (We consider data starting eight years before the default and three years after the default. The exception is Russia. The data for Russia starts in April 1992.) (3) average risk *before* the default; (4) average risk *after* the default; (5) percentage of months before the default with risk above the after-default average; (6) percentage of months after the default with risk below the before-default average; (7) number of consecutive months before the default with political risk below the after-default average; (8) number of consecutive months after the default with political risk above the before-default average; n/a indicates that political risk after the default is always above the before-default average.

the default. One can see in Table 1 that this occurs in Argentina, Ecuador, Pakistan, and Russia.

Among the four default episodes associated with an increase in political risk, we identify the default episode in Argentina as the most likely to have been political. Argentina exhibits the largest increase in political risk after the default. Comparing columns (3) and (4) in Table 1, we can see that the post-default level of political risk in Ecuador, Pakistan, and Russia is less than 10 percent higher than the pre-default level. In Argentina, the post-default level of political risk is 39 percent higher than the pre-default level. In addition, among these four countries, Argentina is the most likely to have experienced the kind of political stability Hatchondo, Martinez, and Sapriza (2009) argue should precede a political default. Recall that Hatchondo, Martinez, and Sapriza (2009) explain that pre-default creditor-friendly governments would only choose debt levels for which a political default would occur in environments with high political stability. Table 1 shows that among the four countries where default episodes marked an increase in political risk, Argentina is the only one where the level of political risk was consistently lower before the default (and consistently high after the default; see columns (5)–(8) in Table 1). In order to further support the view that the Argentine default was preceded by political turnover, the next subsection describes political events around the default.

Political Turnover Around the Argentine Default

A series of political events that occurred around the 2001 default seem to confirm that the default episode in Argentina was preceded by political turnover.

In the presidential campaign of 1999, the two main candidates expressed opposing positions as to whether the future government should declare a moratorium on its foreign debt. *The Economist* (1999) wrote that "while Eduardo Duhalde, his Peronist opponent, has made rash public-spending promises, and suggested that Argentina should default on its foreign debt, it has been Mr. de la Rua who has responsibly promised to maintain the main thrust of current economic policies, including convertibility."

The creditor-friendly approach of Fernando de la Rua's government is also apparent from its attempt to impose drastic austerity to balance the budget—including cuts of up to 13 percent in public sector wages and pensions. In the face of a drying up of credit and an economy in its fourth year of recession, this was perceived to be the only way to stave off default on Argentina's $128 billion of public foreign debt and maintain the currency-board system that pegs the peso, at par, to the dollar. This policy stance was reinforced by de la Rua's statement that "... there'll be no default and no devaluation. Our effort is to reactivate the internal market, which needs lower interest rates. It could be necessary to lower the costs of the debt, but we will comply with our obligations" (see *The Economist* [2001]).

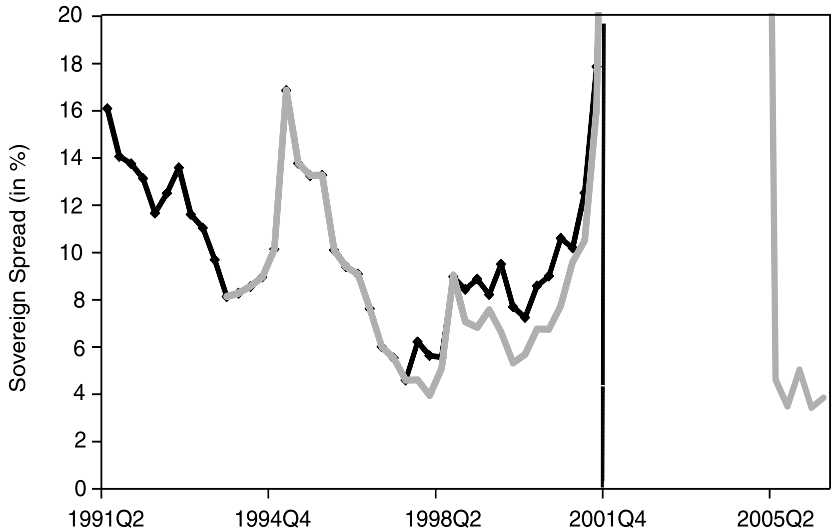
Having lost political support even from members of his own party, de la Rua left office on December 19, 2001, and was succeeded by governments with a more debtor-friendly approach. The newly appointed president, the Peronist Adolfo Rodriguez Saa, immediately declared a default that was widely supported in Congress. He was replaced two weeks later and his successor, Eduardo Duhalde, confirmed the default decision by failing to serve a USD 28 million interest payment due on an Italian lira bond. According to Sturzenegger and Zettelmeyer (2006), it is estimated that around 60 percent of the debt in default was held by domestic residents.

The Behavior of Spread and Debt Levels Around the Argentine Default

Hatchondo, Martinez, and Sapriza (2009) predict that in a political default, post-default debt levels are lower than pre-default levels and post-default spread levels are lower than pre-default levels. We will contrast these predictions with the behavior of debt and spread levels around the 2001 Argentine default, which we have argued has the characteristics of a political default.

Figure 3 shows that, in Argentina, spreads were lower after the 2005 debt exchange, when (according the ICRG index of political risk) the government was perceived as riskier to creditors, than before the default, when the government was perceived as less risky to creditors. Thus, the behavior of the

Figure 3 Argentina Sovereign Spread



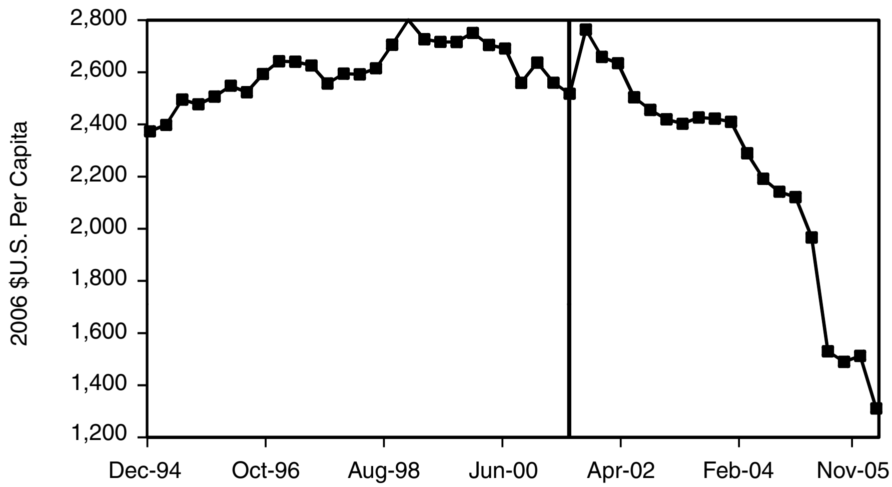
Notes: The vertical line marks the month of default. The line with dots (black) corresponds to the yield of government debt computed by Neumeyer and Perri (2005). The solid gray line corresponds to the measure of the spread computed by J.P. Morgan using foreign currency debt.

spread in Argentina is roughly in line with the one predicted by Hatchondo, Martinez, and Sapriza (2009).[8]

Figure 4 shows that in Argentina, governments perceived to be riskier to creditors have chosen relatively low debt levels after the default—the debt level decreases sharply in 2005 when the defaulted debt is exchanged. This is consistent with the decrease in the debt level after a political default predicted by Hatchondo, Martinez, and Sapriza (2009). It is also consistent with the difficulties in market access observed after a default episode (IMF [2002a] and Gelos, Sahay, and Sandleris [2004] discuss evidence of a drainage in capital flows to countries that defaulted).

[8] In Hatchondo, Martinez, and Sapriza's (2009) model, the recovery rate on defaulted bonds is zero and, consequently, defaulted bonds have no value. Therefore, Hatchondo, Martinez, and Sapriza (2009) do not present predictions that one could contrast with the spread data between the default episode in 2001 and the debt exchange in 2005.

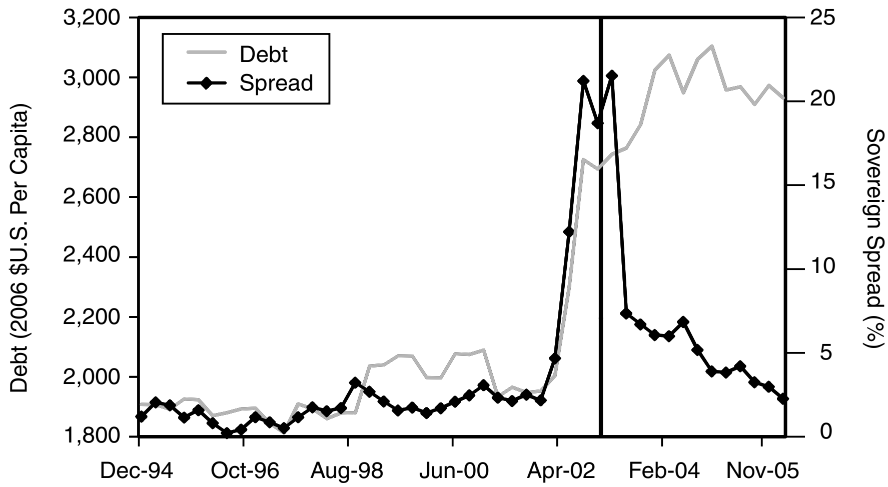
Figure 4 Face Value of Argentina's Public Debt that is Denominated in Foreign Currency



Notes: The series does not include arrears. The vertical line marks the month of default.

Of course, other factors besides political turnover may have affected Argentina's borrowing decisions and the market price of its debt. One way of controlling for some of these factors is to compare the behavior of debt and spread in Argentina with the one in Uruguay. Argentina and Uruguay are neighboring countries with highly correlated business cycles. In fact, both countries had experienced negative growth since 1999, after the Brazilian devaluation. Brazil was a major trading partner of Argentina and Uruguay and both countries had pegged their exchange rate to the dollar, which may have slowed down the adjustment of prices to that shock. Both countries defaulted on their debt, but the 2003 Uruguayan default does not seem to have been triggered by political turnover. According to Table 1, the pre- and post-default levels of political risk in Uruguay are almost identical. There is also anecdotal evidence consistent with that. The Uruguayan president at that time, Jorge Batlle, had previously campaigned in 1989 with a platform that proposed to swap the central banks' gold reserves to pay off the debt in default. In the midst of the 2002 crisis, he announced that the country would make sacrifices in order to honor its debt contracts. Unlike in Argentina, the ruling coalition in Uruguay had control of Congress and managed to approve several rounds of spending cuts and tax increases to reduce the budget deficit (see *The Economist* [2002]). The Uruguayan government could avoid missing

Figure 5 Uruguay's Foreign-Currency-Denominated Public Debt and Sovereign Spread over U.S. Treasury Bills



Notes: The vertical line marks the month of default.

debt payments and also stop a bank run thanks to a joint rescue package provided by the IMF, the World Bank, and the Inter-American Development Bank (see IMF [2002c]). In a press release, the IMF executive board "...commended the Uruguayan authorities for their decisive policy action, their commitment to maintaining a framework that will foster private sector activity, and their continued close cooperation with the Fund..." (see IMF [2002b]). Sturzenegger and Zettelmeyer (2006) estimate that the bondholders that participated in the Uruguayan exchange suffered a reduction in the net present value of their claims within the range of 10–15 percent, substantially lower than the loss experienced by holders of Argentine debt (more than 60 percent). In order to induce a higher participation rate in their debt exchange, the Uruguayan authorities announced that the new bonds were going to receive de facto seniority over the previously issued bonds. Ex post, bondholders that did not participate in the exchange were fully paid back.

Figure 5 shows that the spread and debt levels in Uruguay were not lower after the default episode than before the crisis (as they were in Argentina). The figure also shows that the spread and debt levels were not particularly low in Uruguay after 2005, at the time when they were low in Argentina. Thus, we do not find that low post-default levels of spread and debt in Argentina may be accounted for by shocks that also affected Uruguay during that time.

4. CONCLUSIONS

This article discusses how political factors may influence sovereign default risk. First, the article presents a summary of theoretical studies on this issue. We survey studies that argue that a sovereign may be willing to repay its debt because it is in the best interest of local agents with political power. We also discuss theoretical studies that examine how changes in the government's willingness to pay and the frequency of these changes (political stability) affect sovereign default risk. We then discuss a large body of empirical work that finds evidence of the influence of political stability and other characteristics of a political system on default risk. In addition, we study five recent sovereign defaults and find that the 2001 Argentine default is the most likely to have been triggered by political turnover, and that the behavior of spread and debt levels around that default is broadly in line with the one predicted by Hatchondo, Martinez, and Sapriza (2009).

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